Identifying preference for early resolution from asset prices

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This paper develops an asset market based test for preference for the timing of resolution of uncertainty. Our main theorem provides a characterization of preference for early resolution of uncertainty in terms of the risk premium of assets realized during the period when the informativeness of macroeconomic announcements is resolved. Empirically, we find support for preference for early resolution of uncertainty based on evidence on the dynamics of the implied volatility of S&P 500 index options before FOMC announcements.

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Key words: preference for early resolution of uncertainty, generalized risk sensitivity, macroeconomic announcements, volatility

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1 Introduction

In this paper, we develop a revealed preference theory that allows us to use asset-market-based evidence to detect investors' preference for the timing of resolution of uncertainty. Our main theorem states that the representative agent prefers early (late) resolution of uncertainty if and only if claims to market volatility, which can be constructed from index options, require a *positive* (negative) premium during the period when the informativeness of macroeconomic announcements is resolved. Empirically, using evidence on the implied volatility of S&P 500 index options around Federal Open Market Committee (FOMC) announcements, we find supportive evidence for investors' preference for *early* resolution of uncertainty.

The notion of preference for the timing of resolution of uncertainty is formally developed in Kreps and Porteus [32]. Models with preference for early resolution (PER) of uncertainty—in particular, the recursive preference with constant elasticity—have been widely applied in the asset pricing literature (see, e.g., Epstein and Zin [17, 19], Weil [48], Bansal and Yaron [5], and Hansen, Heaton, and Li [23], among others). However, in the constant elasticity recursive utility model and in most applied asset pricing models, PER is typically intertwined with other aspects of preferences, such as risk aversion and intertemporal elasticity of substitution (IES). As a result, the exact role for PER in asset pricing is not well understood. In addition, the asset pricing implications of models with PER are typically similar to a broad class of preferences that satisfy generalized risk sensitivity (Ai and Bansal [2]). The purpose of this paper is to provide an equivalent characterization of PER in terms of asset prices and use asset market data to identify investors' preference for the timing of resolution of uncertainty.

Preferences are often the starting point of macroeconomic analysis and asset pricing studies. Modern economic theory implies that asset prices are evaluated using marginal utilities. Therefore, the empirical evidence from asset markets can potentially provide valuable guidance for the choice of preferences in macroeconomic analysis in general and in policy studies in particular. However, results that allow researchers to use relevant asset-market-based evidence to identify exact properties of preferences are rare. In this paper, we provide a general result that allows researchers to build such links and apply our result to establish a necessary and sufficient condition for PER in terms of asset prices. We show that the representative investor prefers early resolution of uncertainty if and only if claims to market volatility require a positive premium during the period of *resolution of information quality (ROIQ)*, that is, the period during which the uncertainty about the informativeness of macroeconomic announcements is resolved. We provide empirical evidence for investors' preference for the timing of resolution of uncertainty based on our theoretical insights and find evidence supportive of PER.

Our main theorem builds on the notion of generalized risk sensitivity (GRS) developed in Ai and Bansal [2]. Ai and Bansal [2] define GRS to be the class of all preferences in which the marginal utility of consumption decreases with respect to continuation utility. The theorem of generalized risk sensitivity in Ai and Bansal [2] demonstrates that a non-negative announcement premium for all assets that are comonotone with continuation utility is equivalent to GRS. However, GRS is a very general condition that includes many examples of non-expected utility as special cases—for example, the Gilboa and Schmeidler [20] maxmin expected utility, which is indifferent between an early or late resolution of uncertainty, and the Kreps and Porteus [32] utility, which prefers early resolution of uncertainty. The announcement premium itself does not allow us to identify PER.

The condition of GRS, however, implies that the ranking of the marginal utility of consumption is the inverse ranking of the level of continuation utility and allows us to design a thought experiment to identify PER from risk premia. PER implies that the utility level of the representative agent is higher when she expects a more informative macroeconomic announcement and lower when she expects a non-informative announcement. The key insight of our paper is that under GRS, PER is equivalent to a negative comonotonicity between marginal utility and the expected informativeness of the upcoming macroeconomic announcement. Because more informative macroeconomic announcements are associated with higher realized stock market volatility upon announcements, the risk premium on claims to market volatility can be used to detect the ranking of marginal utility with respect to the informativeness of macroeconomic announcements and, therefore, PER. The asset pricing test implied by our theorem is easily implementable as claims to market volatility can be replicated using a portfolio of options.

Based on the above insight, we design an empirical exercise to identify PER from asset market data. Our empirical exercise contains two steps. The first step is to identify a period of resolution of the information quality (ROIQ) of FOMC announcements, which is when investors learn whether or not the upcoming announcement is informative. The second step is to estimate the risk premium for claims to market volatility associated with FOMC announcements to identify PER. Based on standard results from option pricing—for example, Carr and Madan [11], Britten-Jones and Neuberger [9], Bakshi, Kapadia, and Madan [3], and Jiang and Tian [27]—we construct a replicating portfolio for market volatility and find evidence of a positive premium, which is consistent with preference for early resolution of uncertainty. We find that standard straddle-based volatility trades also embed this positive premium.

Related literature Our theoretical work builds on the literature that studies decision making under non-expected utility. We adopt the general representation of dynamic preferences of Strzalecki [46]. The generality of our approach is important given that our purpose is to identify the property of preferences from asset market data and given that PER is often intertwined with other aspects of preferences in the popular recursive utility formulation used in applied asset pricing work.[1] In particular, the general setup allows us to distinguish different decision-theoretic concepts such as generalized risk sensitivity, uncertainty aversion, and preference for early resolution of uncertainty.

Our framework includes most of the non-expected utility models in the literature as special cases, such as the maxmin expected utility of Gilboa and Schmeidler [20], the dynamic version of which is studied by Chen and Epstein [12] and Epstein and Schneider [15]; the recursive preference of Kreps and Porteus [32] and Epstein and Zin [17]; the robust control preference of Hansen and Sargent [25, 26] and the related multiplier preference of Strzalecki [45]; the variational ambiguity-averse preference of Maccheroni, Marinacci, and Rustichini [35, 36]; the smooth ambiguity model of Klibanoff, Marinacci, and Mukerji [30, 31]; and the disappointment aversion preference of Gul [22].

Earlier work on the revealed preference approach for expected utility includes Green and Srivastava [21] and Epstein [18]. More recently, Kubler, Selden, and Wei [33] and Echenique and Saito [13] develop asset-market-based characterizations of the expected utility model. Unlike our paper, none of these papers focus on GRS and aim to connect their results to asset market data.

Our paper is related to several papers that study PER in asset pricing models. Ai [1] demonstrates that in a production economy with long-run risk, most of the welfare gain from knowing more information about the future is a result of PER, not because agents can use the information to improve the intertemporal allocation of resources. Epstein, Farhi, and Strzalecki [14] show that in the calibrated long-run risk model, the representative agent is willing to pay more than 30% of her permanent income to resolve all future uncertainty, and they argue that this magnitude is implausibly high by introspection. They also state, "*We are not aware of any market-based or experimental evidence that might help with a quantitative assessment,*" whereas our paper provides market-based empirical evidence for PER without assuming a parametric utility function. Kadan and Manela [29] estimate the value of information in a model with recursive utility. Schlag, Thimme, and Weber [43] find suggestive evidence for PER using options market data. Both of these papers assume the

[1] For example, in the constant elasticity case, as shown in Ai and Bansal [2], PER is equivalent to risk aversion being higher than IES, which is also equivalent to GRS.

CES form of utility function and do not distinguish PER from GRS, or uncertainty aversion.

A vast literature applies the above non-expected utility models to the study of asset prices and the equity premium. We refer readers to Epstein and Schneider [16] for a review of asset pricing studies with the maxmin expected utility model, Ju and Miao [28] for an application of the smooth ambiguity-averse preference, Hansen and Sargent [24] for the robust control preference, Routledge and Zin [41] for an asset pricing model with disappointment aversion, and Bansal and Yaron [5], Bansal [4], and Hansen, Heaton, and Li [23] for the long-run risk model that builds on recursive preferences. Borovicka and Stachurski [8] provide necessary and sufficient conditions for the existence and uniqueness of recursive preferences with constant elasticities. Bhamara and Uppal [6] study the role of risk aversion and intertemporal elasticity of substitution in portfolio choice problems. Bidder and Dew-Becker [7] link ambiguity aversion to long-run risk models. Skiadas [44] provides an excellent textbook treatment of recursive-preferences-based asset pricing theory.

Our paper is also related to the previous research on stock market returns on macroeconomic announcement days. The previous literature documents that stock market returns and Sharpe ratios are significantly higher on days with macroeconomic news releases both in the United States (Savor and Wilson [42]) and internationally (Brusa, Savor, and Wilson [10]). Lucca and Moench [34] find similar patterns and document a pre-FOMC announcement drift. Mueller, Tahbaz-Salehi, and Vedolin [38] document an FOMC announcement premium on the foreign exchange market and attribute it to compensation to financially constrained intermediaries.

The rest of the paper is organized as follows. We begin with a simple example in Section 2 to illustrate the concept of preference for early resolution of uncertainty and generalized risk sensitivity. In Section 3, we develop a thought experiment that allows us to identify PER from the risk premia of claims to market volatility. Building on these theoretical insights, in Section 4, we develop an identification strategy and present evidence for PER based on option prices on S&P 500 index options. Section 5 concludes.

2 PER and GRS

In this section, we illustrate the concepts of preference for early resolution of uncertainty and generalized risk sensitivity in a simple three-period model. We also provide simple examples for both properties of preferences. To set up some notation, we consider an economy with three periods, 0, 1, 2. Let (S, Σ, μ) be a finite probability space with equal probabilities.

We denote $S = \{1, 2, \cdots n\}$, where $\mu(s) = \frac{1}{n}$ for $s = 1, 2, \cdots$. Let $(\Omega, \mathcal{F}, \boldsymbol{\mu}) = (S, \Sigma, \mu)^3$ be the product space, and let $\mathcal{L}(\Omega, \mathcal{F}, \boldsymbol{\mu})$ be the set of real-valued random variables. A typical realization of states is denoted as (s_0, s_1, s_2), where for $t = 0, 1, 2$, s_t is the realization of the state in period t. A consumption plan is denoted as $C = [c_0(s_0), c_1(s_0, s_1), c_2(s_0, s_1, s_2)]$, where consumption in each period is a measurable function of history: $c_t : (S, \Sigma)^t \to \boldsymbol{C}$, for $t = 0, 1, 2$. Here, the feasible set of consumption, $\boldsymbol{C}$ is a subset of the positive orthant of the real line $\boldsymbol{R}$. To simplify notation, we use the convention that $s^t = \{s_0, s_1, \cdots s_t\}$ denotes the history of s up to time t.

As in Ai and Bansal [2], we consider conditional preferences induced by a triple $\{u, \beta, \mathcal{I}\}$, where $u : \boldsymbol{C} \to \boldsymbol{R}$ maps consumption into utility units, β is the discount rate, and $\mathcal{I} : \mathcal{L}(\Omega, \mathcal{F}, \boldsymbol{\mu}) \to \boldsymbol{R}$ is a certainty equivalent functional that maps continuation utility, which is a random variable, into the real line. In our setup, date-t utility is constructed recursively using

$$V_t(C)\left(s^t\right) = u\left(c_t\left(s^t\right)\right) + \beta\mathcal{I}\left[V_{t+1}(C)\left(s^{t+1}\right)\right], \tag{1}$$

for $t = 0, 1$, where the terminal utility on date 2 is given by $V_2(C) = u(c_2)$. Here, we use the notation $V_t(C)(s^t)$ for the date-t utility of the consumption plan C at history s^t.[2] To simplify notation, we will suppress C and simply write $V_t(s^t)$ whenever the underlying consumption plan is clear from the context.

2.1 Preference for early resolution of uncertainty

To provide a definition of preference for early resolution of uncertainty, we restrict our attention to a simple class of consumption plans. We consider two consumption plans, $C^E = [\bar{c}_0, \bar{c}_1, c_2(s_1)]$ and $C^L = [\bar{c}_0, \bar{c}_1, c_2(s_2)]$, where both $\bar{c}_0$ and $\bar{c}_1$ are constants, and $c_2 : (S, \Sigma, \mu) \to \boldsymbol{C}$ is a random variable that depends only on s but not its history. Note that both plans, C^E and C^L, have the same unconditional distribution because s_1 and s_2 do. However, under C^E, which represents early resolution, period-2 consumption, $c_2(s_1)$, is known in period 1 because s_1 is realized in period 1. By contrast, under C^L, which represents late resolution, the uncertainty in s_2 only realizes in period 2.

A dynamic preference represented by $\{u, \beta, \mathcal{I}\}$ is said to satisfy preference for early resolution of uncertainty if $V_0(\bar{c}_0, \bar{c}_1, c_2(s_1)) \geq V_0(\bar{c}_0, \bar{c}_1, c_2(s_2))$ for all $\bar{c}_0$, $\bar{c}_1$, and all measurable functions $c_2(s)$. Our concept of PER is the same as Kreps and Porteus [32].

[2] Strictly speaking, to emphasize the dependence of $\mathcal{I}_t$ on period-t information, we should allow $\mathcal{I}_t : L(\Omega, \mathcal{F}, \boldsymbol{\mu}) \to \boldsymbol{R}$ to be a family of certainty equivalent functionals indexed by t. For each t, $\mathcal{I}_t$ maps $(S, \Sigma)^{t+2}$ measurable functions into $(S, \Sigma)^{t+1}$ measurable functions.

Figure 1 provides a graphic illustration of C^E (top panel) and C^L (bottom panel) for the case in which $c_2(s)$ takes on only two values, c_U and c_D, where $c_U > c_D$. The squares represent the consumption in each period, and the circles represent the agent's information node. The uncertainty is resolved early in period 1 in the top panel under consumption plan C^E, where the agent is able to distinguish nodes 1_U and 1_D. The bottom panel illustrates late resolution of uncertainty under consumption plan C^L, where the value of $c_2(s_2)$ is known only in period 2.

Figure 1: **Early and late resolution of uncertainty**

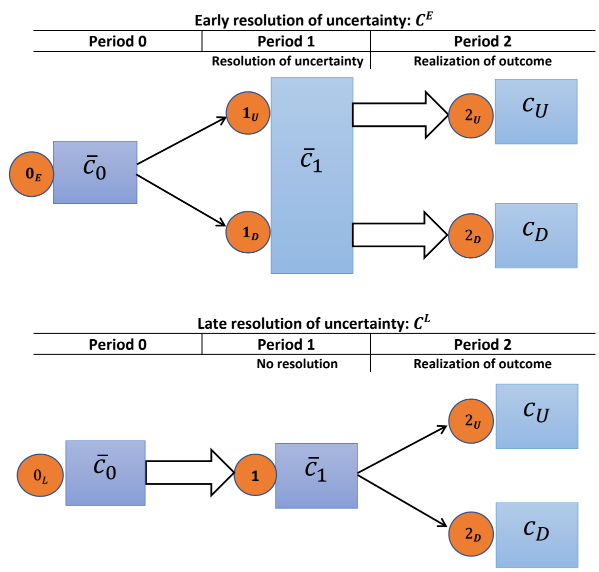


Figure 1 illustrates the notion of PER. Both panels have identical unconditional distributions of consumption. The top panel is a situation with early resolution, as the uncertainty about $c_2(s_1)$ is resolved one period earlier in period 1. The bottom panel corresponds to the case of late resolution because the value of $c_2(s_2)$ is not revealed to the consumer until period 2.

Recursion (1) allows us to compute the utility associated with C^E and C^L. Under C^E, there is no uncertainty at the end of period 1 because period-2 consumption is perfectly predictable. Therefore, period-1 utility is computed as $V_1\left(C^E\right)(s_1) = u(\bar{c}_1) + \beta u(c_2(s_1))$. The time-0 utility is given by

$$V_0\left(C^E\right) = u(\bar{c}_0) + \beta \mathcal{I}\left[u(\bar{c}_1) + \beta u(c_2(s_1))\right]. \tag{2}$$

In the case of late resolution (bottom panel of Figure 1), because uncertainty is resolved in period 2, we first need to aggregate over uncertain states of the world when computing period-1 utility $V_1\left(C^L\right) = u(\bar{c}_1) + \beta \mathcal{I}\left[u(c_2(s_2))\right]$, and simply aggregate over time in period

0 to get

$$V_0\left(C^L\right) = u\left(\bar{c}_0\right) + \beta\left\{u\left(\bar{c}_1\right) + \beta\mathcal{I}\left[u\left(c_2\left(s_2\right)\right)\right]\right\}. \tag{3}$$

A comparison of equations (2) and (3) makes it clear that PER can be formulated as the following property of the certainty equivalent functional:

$$\mathcal{I}\left[u\left(\bar{c}_1\right) + \beta u\left(c_2\left(s_1\right)\right)\right] \geq u\left(\bar{c}_1\right) + \beta\mathcal{I}\left[u\left(c_2\left(s_2\right)\right)\right]. \tag{4}$$

Below we provide simple examples of recursive preference that may satisfy preference for early or late resolution of uncertainty depending on the value of the discount factor.

Examples In this section, we compute the utility level at node 0_E for the case of early resolution of uncertainty and that at node 0_L for the case of late resolution of uncertainty for several examples of preferences. Our first example is the expected utility, where $\mathcal{I}\left(u\right) = E\left[u\right]$. Expected utility is indifferent toward the timing of resolution of uncertainty. The utility associated with early resolution,

$$V_0\left(0_E\right) = u\left(\bar{c}_0\right) + \beta E\left[u\left(\bar{c}_1\right) + \beta u\left(c_2\left(s_1\right)\right)\right] = u\left(\bar{c}_0\right) + \beta u\left(\bar{c}_1\right) + \beta^2 E\left[u\left(c_2\left(s_1\right)\right)\right],$$

and that associated with late resolution,

$$V_0\left(0_L\right) = u\left(\bar{c}_0\right) + \beta\left\{u\left(\bar{c}_1\right) + \beta\mathcal{I}\left[u\left(c_2\left(s_2\right)\right)\right]\right\} = u\left(\bar{c}_0\right) + \beta u\left(\bar{c}_1\right) + \beta^2 E\left[u\left(c_2\left(s_2\right)\right)\right],$$

are the same.

Our second example is the multiple-prior expected utility of Gilboa and Schmeidler [20] and Chen and Epstein [12]. We assume that $\mathcal{I}\left(u\right) = \min_{\phi\in\Phi} E\left[\phi u\right]$, where Φ is a set of probability densities. We assume that the $\mathcal{I}$ operator defined by Φ is distribution invariant. That is, for any u and v, if u and v have the same probability distribution under P, then $\min_{\phi\in\Phi} E\left[\phi u\right] = \min_{\phi\in\Phi} E\left[\phi v\right]$. The utility for early resolution at node 0_E is

$$V_0\left(0_E\right) = u\left(\bar{c}_0\right) + \beta\min_{\phi\in\Phi} E\left[\phi\left\{u\left(\bar{c}_1\right) + \beta u\left(c_2\left(s_1\right)\right)\right\}\right] = u\left(\bar{c}_0\right) + \beta u\left(\bar{c}_1\right) + \beta^2 E_{\phi\in\Phi}\left[\phi u\left(c_2\left(s_1\right)\right)\right].$$

The utility associated with later resolution of uncertainty at node 0_L can be computed as

$$V_0\left(0_L\right) = u\left(\bar{c}_0\right) + \beta\left\{u\left(\bar{c}_1\right) + \beta\min_{\phi\in\Phi} E\left[u\left(c_2\left(s_2\right)\right)\right]\right\} = V_0\left(0_E\right),$$

where the last equality holds because $c_2\left(s_1\right)$ and $c_2\left(s_2\right)$ have the same distribution.

Our third example is the multiplier robust control preference of Hansen and Sargent [24]. Here we assume that $u : \boldsymbol{C} \to \boldsymbol{R}$ is strictly increasing and $\mathcal{I}(u) = -\theta \ln E\left[e^{-\frac{1}{\theta}u}\right]$ for some parameter $\theta > 0$. In the appendix, we show that this choice of the aggregator has preference for early (late) resolution if $\beta < (>)\, 1$ and is indifferent toward the timing of resolution of uncertainty if $\beta = 1$. To simplify computation, we assume that $u(c) = \ln(c)$ and $\ln c_2(s) \sim N(\mu, \sigma^2)$. To calculate the utility associated with early resolution,

$$\begin{aligned} V_0(0_E) &= \ln \bar{c}_0 + \beta \mathcal{I}\left[\ln \bar{c}_1 + \beta \ln c_2(s_1)\right] && (5) \\ &= \ln \bar{c}_0 - \beta\theta \ln E\left[e^{-\frac{1}{\theta}[\ln \bar{c}_1 + \beta \ln c_2(s_1)]}\right] && (6) \\ &= \ln \bar{c}_0 + \beta \ln \bar{c}_1 + \beta^2 \mu - \frac{1}{2\theta}\beta^3\sigma^2. && (7) \end{aligned}$$

The utility associated with late resolution is

$$\begin{aligned} V_0(0_L) &= \ln \bar{c}_0 + \beta \ln \bar{c}_1 + \beta^2 \mathcal{I}\left[\ln c_2(s_2)\right] && (8) \\ &= \ln \bar{c}_0 + \beta \ln \bar{c}_1 - \beta^2\theta \ln E\left[e^{-\frac{1}{\theta}\ln c_2(s_2)}\right] && (9) \\ &= \ln \bar{c}_0 + \beta \ln \bar{c}_1 + \beta^2 \mu - \frac{1}{2\theta}\beta^2\sigma^2. && (10) \end{aligned}$$

Clearly, if $\beta < 1$, then $V_0(0_E) > V_0(0_L)$, and the above specified aggregator has preference for early resolution. In fact, under the assumption of $u(c) = \ln c$ and $\beta < 1$, we have $V(t) = \ln c_t - \theta\beta \ln E\left[e^{-\frac{1}{\theta}V(t+1)}\right]$. This is recognized as the Epstein-Zin preference with unit IES and a risk aversion of $1 + \theta$. It is well known that the Epstein-Zin preference has preference for early resolution if risk aversion is higher than the inverse of IES. In our example, this condition is guaranteed by $\theta > 0$.

If we assume $\beta > 1$, then $V_0(0_E) < V_0(0_L)$, and the resulting preference has preference for late resolution of uncertainty. The case $\beta > 1$ is typically not discussed in the literature, but as we will see in the following section, the case $\beta > 1$ provides a convenient example that has preference for late resolution of uncertainty and at the same time, satisfies generalized risk sensitivity.

From a decision theory perspective, Kreps and Porteus [32] note that preference for early resolution of uncertainty can be motivated either by pure preference or by un-modeled planning. Epstein, Farhi, and Strzalecki [14] develop a thought experiment and compute how much the representative agent is willing to pay to resolve all future uncertainty in long-run risk models. In our setup, as a result of indifference toward the timing of resolution of uncertainty, an expected utility maximizer and a multiple-prior expected utility maximizer are not willing to pay anything in exchange for information that they cannot act upon. Due to preference for early resolution, an agent with the multiplier robust control preference with $\beta < 1$ is willing to pay a positive amount for information about future consumption. Epstein, Farhi, and Strzalecki [14] remark that there is no asset-market-based evidence to infer the consumer's preference for early resolution of uncertainty. The purpose of this paper is to provide such evidence.

As shown in Strzalecki [46], general characterizations of property (4) in terms of the functional form of $\mathcal{I}$ can be quite complicated. Directly testing the functional form of $\mathcal{I}$ from asset prices seems to be extremely hard. The asset pricing test we propose in this paper takes advantage of the notion of generalized risk sensitivity developed in Ai and Bansal [2], which we briefly review in the following section.

2.2 Generalized risk sensitivity

To discuss the notion of generalized risk sensitivity, we first introduce some terminology. Let $X : (\Omega, \mathcal{F}, P) \to R$ and $Y : (\Omega, \mathcal{F}, P) \to R$ be two random variables. Variable X is said to first-order stochastically dominate Y if $E[\phi(X)] \geq E[\phi(X)]$ whenever ϕ is increasing, which we denote as $X \succeq_{FSD} Y$. X strictly first-order stochastically dominates Y if $X \succeq_{FSD} Y$ and if $E[\phi(X)] > E[\phi(X)]$ whenever ϕ is strictly increasing, which we denote as $X \succ_{FSD} Y$. X is said to second-order stochastically dominate Y if $E[\phi(X)] \geq E[\phi(X)]$ whenever ϕ is concave, which we denote as $X \succeq_{SSD} Y$. X strictly second-order stochastically dominates Y if $X \succeq_{SSD} Y$ and $E[\phi(X)] > E[\phi(X)]$ whenever ϕ is strictly concave.[3] We denote strict second-order stochastic dominance as $X \succ_{SSD} Y$. In what follows, we will assume that $\mathcal{I}$ is strictly increasing in first-order stochastic dominance. That is, $\mathcal{I}[X] \geq \mathcal{I}[Y]$ if $X \succeq_{FSD} Y$, and the inequality is strict if $X \succ_{FSD} Y$. This assumption is a requirement of the monotonicity of the preference.

An intertemporal preference represented by $\{u, \beta, \mathcal{I}\}$ is said to satisfy generalized risk sensitivity if $\mathcal{I}$ is increasing in second-order stochastic dominance (see (Ai and Bansal [2])); that is, $\mathcal{I}[X] \geq \mathcal{I}[Y]$ if $X \succeq_{SSD} Y$. It satisfies strict generalized risk sensitivity if $\mathcal{I}$ is strictly increasing in second-order stochastic dominance. Ai and Bansal [2] demonstrate that generalized risk sensitivity provides a necessary and sufficient condition for the existence of announcement premia in representative agent economies.

To illustrate the concept of GRS, we consider the top panel of Figure 1 and interpret the

[3] For other equivalent definitions of second-order stochastic dominance, see Rothschild and Stiglitz [40].

event in period 1 that reveals the true value of $c_2(s_1)$ as an announcement. The utility of the agent at time 0 can be computed in two steps:

$$V_0 = u(\bar{c}_0) + \beta\mathcal{I}[V_1(s_1)],$$

where $\forall s_1 \in \Omega$, the continuation utility $V_1(s_1)$ is computed as

$$V_1(s_1) = u(\bar{c}_1) + \beta u(c_2(s_1)). \tag{11}$$

We compute the stochastic discount factor that prices the period-1 state-contingent payoff into period-0 consumption units, denoted $SDF(s_0, s_1)$. As in standard equilibrium models, the stochastic discount factor can be constructed as the ratio of marginal utilities. Therefore, if we interpret $V_1 = [V_1(1), V_1(1), \cdots V_1(n)]$ as a finite-dimensional vector and denote $\frac{\partial\mathcal{I}[V_1]}{\partial V_1(s_1)}$ as the partial derivative of the certainty equivalent with respect to $V_1(s_1)$,

$$SDF(s_0, s_1) = \beta\frac{\frac{1}{\boldsymbol{\mu}(s_1)}\frac{\partial\mathcal{I}[V_1]}{\partial V_1(s_1)}u'(\bar{c}_1)}{u'(\bar{c}_0)} \propto \frac{\partial\mathcal{I}[V_1]}{\partial V_1(s_1)},$$

where in the last step, we suppress the term $\beta\frac{u'(\bar{c}_1)}{u'(\bar{c}_0)}$, which does not depend on s_1 and does not affect the risk premium. We have also used the fact that $\boldsymbol{\mu}(s) = \frac{1}{n}$ does not depend on s. Clearly, if $\frac{\partial\mathcal{I}[V_1]}{\partial V_1(s_1)}$ is a decreasing function of $V_1(s_1)$, then any payoff that is positively correlated with the continuation utility, $V_1(s_1)$, will require a positive risk premium at the announcement.

Formally, we consider an endowment economy in which aggregate consumption is of the form $C = [\bar{c}_0, \bar{c}_1, c_2(s_1)]$. We think of period 1 as the macroeconomic announcement period during which the value of $c_2(s_1)$ is revealed. We consider a state-contingent payoff $X(s_1)$ and denote the present value of $X(s_1)$ from the perspective of period 0 as $P_0(X)$. We say that asset X provides an announcement premium if $\frac{E[X]}{P_0(X)} > R_f(0)$, where $R_f(0)$ is the risk-free rate between period 0 and period 1.

To establish a link between generalized risk sensitivity and the announcement premium, for any two random variables X and Y, we define X and Y to be comonotone with respect to each other if $\forall s$ and s' such that $X(s) \cdot X(s') \neq 0$,

$$[X(s) - X(s')][Y(s) - Y(s')] \geq 0, \tag{12}$$

and define X and Y to be negatively comonotone with respect to each other if (12) holds with $\leq$. Strict comonotonicity is defined similarly with condition (12) holding with strict inequality. The following theorem is an extension of the theorem of generalized risk sensitivity in Ai and Bansal [2].

Theorem 1. *(Theorem of generalized risk sensitivity) Assuming that both u and $\mathcal{I}$ are strictly increasing and continuously differentiable, the following statements are equivalent:*

1. *The announcement premium for any asset strictly comonotone with $c_2(s_1)$ is strictly positive.*

2. *The certainty equivalent functional, $\mathcal{I}$, is strictly increasing in second-order stochastic dominance.*

3. *For any continuation utility, $V : \{\Omega, \mathcal{F}\} \to R$, the vector of partial derivatives of $\mathcal{I}$ with respect to V, $\left\{\frac{\partial \mathcal{I}[V_1]}{\partial V_1(s)}\right\}_{s=1,2,\cdots,n}$, is strictly negatively comonotone with $\{V_1(s)\}_{s=1,2,\cdots,n}$.*

The above theorem complements Theorem 2 of Ai and Bansal [2]. The discrete state setup allows us to establish the equivalence between strict generalized risk sensitivity and a strictly positive announcement premium, which is not covered by Theorem 2 of Ai and Bansal [2] but is important for identifying preference for early resolution of uncertainty.[4] Below we discuss the generalized risk sensitivity property of the examples of preferences discussed in Section 2.1.

Examples The expected utility does not satisfy strict generalized risk sensitivity. Clearly, if $\mathcal{I}$ is the expectation operator, then $\mathcal{I}[u] = \mathcal{I}[v]$ as long as $E[u] = E[v]$ regardless of the second-order stochastic dominance between u and v.

The multiple-prior expected utility satisfies generalized risk sensitivity and satisfies strict generalized risk sensitivity if Φ is not a singleton. In fact, if $u \succeq_{SSD} v$, then $\min_{\phi\in\Phi} E[\phi u] \geq \min_{\phi\in\Phi} E[\phi v]$, and the inequality is strict if $u \succ_{SSD} v$.[5]

The multiplier robust control preference satisfies generalized risk sensitivity. Using the result from Ai and Bansal [2], aggregators of the form $\mathcal{I}(u) = \phi^{-1}(E[\phi(u)])$ satisfy generalized risk sensitivity if ϕ is concave.[6] It follows immediately that the example in the last section, that is, $\mathcal{I}(u) = -\theta \ln E\left[e^{-\frac{1}{\theta}u}\right]$, satisfies generalized risk sensitivity as long as

[4] The theorem in Ai and Bansal [2] focuses on weak inequalities.

[5] See Lemma 2 in Wasserman and Kadane [47].

[6] If $u \succeq_{SSD} v$, then $E[\phi(u)] \geq E[\phi(v)]$. As a result, $\phi^{-1}(E[\phi(u)]) \geq \phi^{-1}(E[\phi(v)])$ because ϕ is strictly monotonic.

$\theta > 0$. As a result, this utility function may exhibit preference for early or late resolution of uncertainty, depending on the value of β, but it always satisfies generalized risk sensitivity.

In the rest of the paper, we will restrict our attention to preferences that satisfy generalized risk sensitivity. The assumption of generalized risk sensitivity is appealing in our setup for two reasons. First, it is motivated by the empirical fact of positive macroeconomic announcement premia. Second, it links the level of utility, which is a property of preference, to marginal utilities, which can be conveniently tested from asset prices. In particular, under the assumption of GRS, the ranking of the level of utility is exactly the reverse of the ranking of continuation utility, a property that we exploit in the following sections.

3 An asset pricing test for PER

3.1 A thought experiment

In this section, we extend the three-period model above to construct a thought experiment in which asset prices can be used to identify preference for early resolution of uncertainty. To do so, we combine the early resolution of uncertainty case and the late resolution of uncertainty case in Figure 1 and add a period -1 to construct a four-period model, as illustrated in Figure 2.

In our four-period model, a general consumption plan is denoted as $C = [c_{-1}, c_0(s_0), c_1(s_0, s_1), c_2(s_0, s_1, s_2)]$. To identify PER, it is enough to restrict attention to the class of consumption plans where $C = \left[\bar{c}_{-1}, \bar{c}_0, \bar{c}_1, c_2\left(s_{\iota(s_0)}\right)\right]$, where as before, $c_2 : (S, \Sigma, \mu) \rightarrow \boldsymbol{C}$ is a random variable taking values in the consumption set $\boldsymbol{C}$, and $\bar{c}_t$ are constants for $t = -1, 0, 1$. In addition, $\iota : (S, \Sigma, \mu) \rightarrow \{1, 2\}$ is a random variable that takes a value of either 1 or 2. As illustrated in the previous example, $\iota(s_0) = 1$ represents the case of early resolution of uncertainty and $\iota(s_0) = 2$ represents the case with late resolution of uncertainty.

As illustrated in Figure 2, early or late resolution is a stochastic outcome to be learned in period 0. We call period 0 the period of *resolution of information quality (ROIQ)*, which resolves the informativeness of the upcoming announcement in period 1. The node 0_E represents a situation with early resolution where $\iota(s_0) = 1$ and the continuation utility of the agent, $V_0(0_E)$, can be calculated as in equation (2). The node 0_L represents a situation of late resolution where $\iota(s_0) = 2$ and $V_0(0_L)$ is calculated as in equation (3). In period -1, before the resolution of information quality, the agent's utility is calculated

Figure 2: **Resolution of information quality**

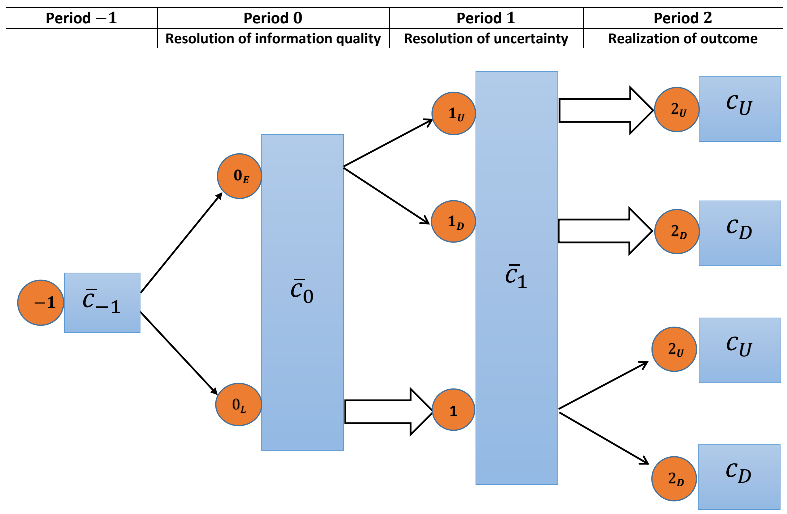


Figure 2 represents our thought experiment of resolution of information quality. The node 0_E is where $\iota(s_0) = 1$ and the agent expects the uncertainty about $c_2(s_1)$ to be resolved in period 1 with an informative macroeconomic announcement that reveals s_1. Node 0_L represents the situation in which $\iota(s_0) = 2$, and therefore the upcoming announcement is expected to be uninformative about $c_2(s_2)$.

as $V_{-1} = u(\bar{c}_{-1}) + \beta \mathcal{I}[V_0(s_0)]$. In our model, the stochastic discount factor that converts period 0 payoff into period -1 consumption units can be calculated as the marginal rate of substitution of consumption between periods 0 and -1:

$$SDF(s_{-1}, s_0) = \frac{\beta \frac{1}{\boldsymbol{\mu}(s_0)} \frac{\partial \mathcal{I}[V_0]}{\partial V_0(s_0)} u'(\bar{c}_0)}{u'(\bar{c}_{-1})} \propto \frac{\partial \mathcal{I}[V_0]}{\partial V_0(s_0)}. \tag{13}$$

By Theorem 1, under the assumption of generalized risk sensitivity, the ranking of the level of utility is the inverse of the ranking of the marginal utilities. That is, for any s_0 and s_0', where s_0 is more informative than s_0', preference for early resolution implies $V(s_0) \geq V(s_0')$. Under GRS, this is true if and only if $\frac{\partial \mathcal{I}[V_0]}{\partial V_0(s_0)} \leq \frac{\partial \mathcal{I}[V_0]}{\partial V_0(s_0')}$. Conversely, preference for late resolution is equivalent to $\frac{\partial \mathcal{I}[V_0]}{\partial V_0(s_0)} \geq \frac{\partial \mathcal{I}[V_0]}{\partial V_0(s_0')}$.

Although the ranking of the level of continuation utility is hard to observe, the ranking of marginal utilities can be detected from the asset market. Suppose we find a payoff X that is increasing in informativeness; that is, $X(s_0) \geq X(s_0')$ whenever s_0 is more informative than s_0'. Then, under PER, $X(s_0)$ will be negatively correlated with $SDF(s_{-1}, s_0)$, and therefore the claim to X will receive a positive risk premium. Conversely, under PLR, $X(s_0)$ will be positively correlated with $SDF(s_{-1}, s_0)$, and therefore the claim to X will receive a negative risk premium. This is the basic intuition of our asset pricing test.

In our model, we have assumed that consumption in periods -1, 0, and 1 is constant and does not depend on the signal. This assumption simplifies our analysis and allows us to prove a theorem that identifies PER from asset prices. Empirically, we interpret the resolution of uncertainty in period 1 as the arrival of macroeconomic announcements and interpret the resolution of information quality as the few days before announcements during which the informativeness of the upcoming announcement becomes known to the public. These events happen at a daily or even hourly frequency, and it is impossible for aggregate consumption to respond at this frequency. Our assumption of constant consumption before period 2 captures this feature of the data, which we use to identify PER.

3.2 An equivalence result

Consider any asset with payoff $X : \Omega \to R$. The payoff X is said to be comonotone with informativeness if for any s_0 and s_0', $[\iota(s_0) - \iota(s_0')][X(s_0) - X(s_0')] \leq 0$; that is, the payoff is higher in the case of early resolution ($\iota(s_0) = 1$) than in the case of late resolution ($\iota(s_0) = 2$). Asset X is said to require a positive resolution of information quality premium if

$$E\left[\frac{X(s_0)}{P_{-1}[X(s_0)]}\right] > R_f(-1),$$

where $R_f(-1)$ is the risk-free interest rate from the end of period -1 to period 0. That is, the ROIQ premium is positive if the strategy of purchasing the asset right before the resolution of information quality and selling it immediately afterward earns an expected return higher than the risk-free interest rate.

Theorem 2. *Assuming that both u and $\mathcal{I}$ are strictly increasing, continuously differentiable, and satisfy strict GRS, the following statements are equivalent:*

1. *The premium for any asset comonotone with informativeness is positive (negative) during the period of ROIQ.*
2. *The certainty equivalent functional, $\mathcal{I}$, satisfies preference for early (late) resolution of uncertainty.*

Given the discussion in the last section, it is straightforward that under GRS, PER implies a positive risk premium for payoffs increasing in the informativeness of the upcoming announcement. The converse of this statement is non-trivial and is the theoretical basis for the identification exercise in this paper. If we have a rich set of assets with payoffs increasing in informativeness, and the risk premium of these assets are positive, then we can safely conclude that the representative agent prefers early resolution of uncertainty.

Figure 3: **Evolution of utility levels and marginal utilities**

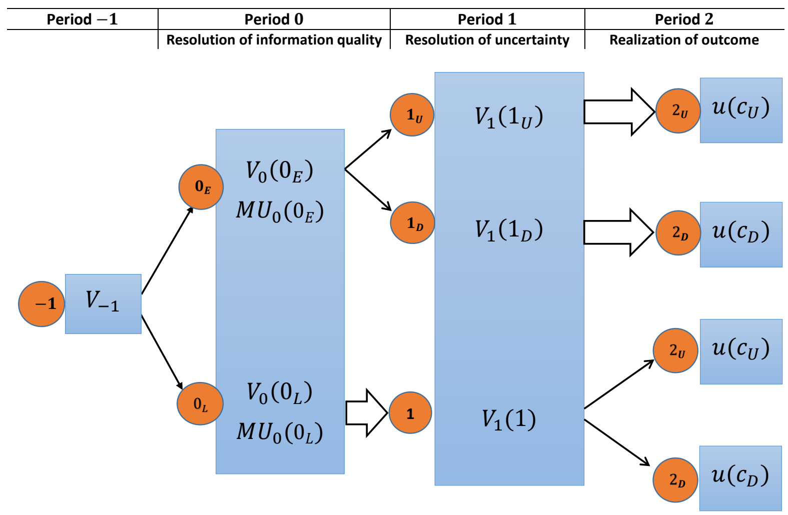


Figure 3 shows the evolution of the utility levels and marginal utilities through our thought experiment. Here, $V_t(s_t)$ and $u_2(c_2)$ are utility levels, and $MU_0(s_0)$ are marginal utilities.

The basic intuition for Theorem 2 can be illustrated using Figure 3, where we mark the level of utility in period 0 and the marginal utility in the same period. The period-2 utilities are simply $u(c_2)$, while those in earlier periods are $V_t(s_t) = u(c_t) + \beta\mathcal{I}\left[V_{t+1}(s_{t+1})\right]$, constructed using the recursive relation. As we explain in Section 2, preference for early resolution implies that $V_0(0_E) > V_0(0_L)$. The ranking of investors' level of utility, however, is typically not observable. Theorem 2 implies that under strict GRS, $V_0(0_E) > V_0(0_L)$ is equivalent to $MU_0(0_E) < MU_0(0_L)$, and the ranking of marginal utility can be detected from asset prices and risk premia. At the same time, if the agent is indifferent toward the timing of resolution of uncertainty, $V_0(0_E) = V_0(0_L)$ and $MU_0(0_E) = MU_0(0_L)$. Because marginal utility is constant in period 0, no asset will receive a risk premium from period -1 to period 0.

Examples We continue with the examples of preferences discussed in Section 2. Under expected utility, $\mathcal{I}\left[V_0\right] = \sum_{s_0}\boldsymbol{\mu}(s_0)V(s_0)$ and $\frac{\partial\mathcal{I}[V_0]}{\partial V_0(s_0)} = \frac{1}{\boldsymbol{\mu}(s_0)}$. Using Equation (13), $SDF(s_{-1}, s_0)$ is equalized across s_0, and there cannot be any preference for early resolution premium.

Under the multiple-prior expected utility, $\mathcal{I}[V_0] = \sum_{s_0} \boldsymbol{\mu}(s_0)\phi^*(s_0)V(s_0)$, where ϕ^* is the minimizing probability density. Because we have already established in Section 2 that the multiple-prior expected utility is indifferent toward the timing of resolution of uncertainty, $V(s_0)$ does not depend on s_0. As a result, any $\phi \in \Phi$ can be used as a minimizing probability. Therefore, $SDF(s_{-1}, s_0)$ is not unique. This is a well-known property for multiple-prior expected utility: it is a concave function, but the set of sub-gradients may not be a singleton. As a result, a positive or a negative preference for early resolution premium can both be consistent with equilibrium.

The multiplier robust control preference is a good example to illustrate the idea of our asset-pricing-based test for PER because it always satisfies GRS but may or may not satisfy PER depending on the value of β. Under the multiplier robust control preference, the SDF can be computed as

$$SDF\left(s_{-1}, s_0\right) = \beta \left(\frac{\bar{c}_0}{\bar{c}_{-1}}\right)^{-1} \frac{e^{-\frac{1}{\theta}V_0(s_0)}}{E\left[e^{-\frac{1}{\theta}V_0(s_0)}\right]}. \tag{14}$$

From the discussion in Section 2.2, generalized risk sensitivity corresponds to $\theta > 0$, in which case $SDF(s_{-1}, s_0)$ is a strictly increasing function of $V_0(s_0)$, as shown in Equation (14). As a result, under generalized risk sensitivity, preference for early resolution of uncertainty (i.e.,$V_0(0_E) > V_0(0_L)$) is equivalent to $SDF(s_{-1}, 0_E) < SDF(s_{-1}, 0_L)$, and preference for late resolution of uncertainty (i.e.,$V_0(0_E) < V_0(0_L)$) is equivalent to $SDF(s_{-1}, 0_E) > SDF(s_{-1}, 0_L)$. Below we show that option prices can be used to identify the ranking of $SDF(s_{-1}, s_0)$ and therefore investors' attitude toward the timing of resolution of uncertainty.

An asset pricing test for PER We continue from the above example and assume that the stock market is a levered claim to aggregate consumption: c_2^{ϕ}, where $\phi > 1$ is a leverage parameter. We first derive the stock market price dynamics. Details on the derivations in this subsection can be found in Appendix B. In period 1, in the case of early resolution of uncertainty, the value of $c_2(s_1)$ is known, and the price of the stock is given by $P_1(s_1) = \beta\bar{c}_1 c_2(s_1)^{\phi}$. In the case of late resolution, the period-1 price is $P_1(s_1) = \beta\bar{c}_1 e^{\phi(\mu+\frac{1}{2}(\phi-\frac{2}{\theta})\sigma^2)}$ and does not depend on the value of c_2. The stock price in period 0 can be calculated accordingly. At the node where the agent expects early resolution, 0_E, $P_0(0_E) = \beta^2\bar{c}_0 e^{\phi(\mu+\frac{1}{2}(\phi-\frac{2\beta}{\theta})\sigma^2)}$, and at the node where the agent expects late resolution, 0_L, $P_0(0_L) = \beta^2\bar{c}_0 e^{\phi(\mu+\frac{1}{2}(\phi-\frac{2}{\theta})\sigma^2)}$.

Note that at node 0_E, the upcoming announcement is expected to be informative. Let $R_f(0_E)$ denote the one-period risk-free interest rate from 0_E to period 1. It is straightforward to see that the announcement premium is positive; that is, $\frac{E[P_1(s_1)]}{P_0(0_E)} = \frac{\beta \bar{c}_1 c_2(s_1)^{\phi}}{\beta^2 \bar{c}_0 e^{\phi(\mu + \frac{1}{2}(\phi - \frac{2\beta}{\theta})\sigma^2)}} = \frac{\bar{c}_1}{\beta \bar{c}_0} e^{\frac{\phi\beta}{\theta}\sigma^2} = R_f(0_E) e^{\frac{\phi\beta}{\theta}\sigma^2} > R_f(0_E)$, as long as $\theta > 0$, regardless of the value of β. That is, regardless of investors' attitude toward the timing of resolution of uncertainty, as long as generalized risk sensitivity is satisfied, the announcement premium is positive. We display the stock market dynamics in Figure 4.

Figure 4: **Evolution of prices of consumption and variance claims**

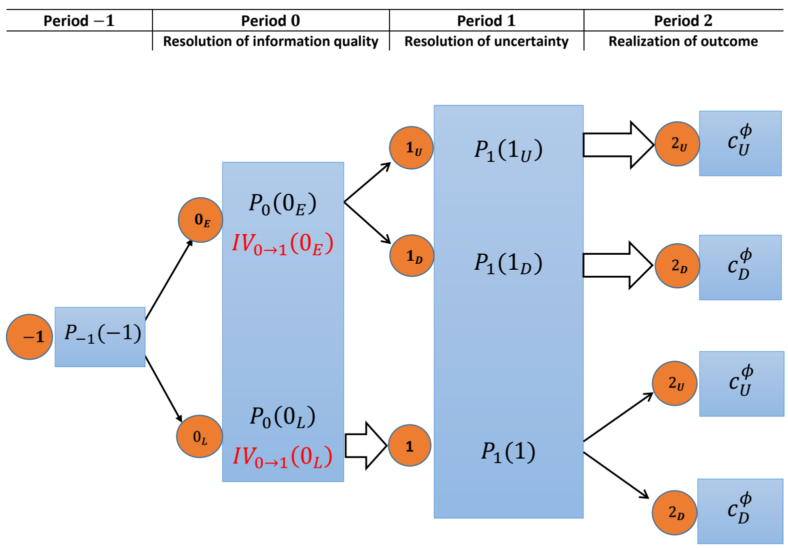


Figure 4 plots the evolution of the prices of the leveraged consumption claim (in black) and the prices of the claim to its period-1 return variance (in red).

To operationalize Theorem 2, we need a test asset with a payoff increasing in informativeness. We show that the option-implied variance can serve as a test asset. Let $IV_{0\to1}(s_0)$ denote the variance of a one-period return from node s_0 to the following node in period 1; that is, $IV_{0\to1}(s_0) = Var[\ln P_1 - \ln P_0 | s_0]$. Here, the subscript $0 \to 1$ indicates the maturity of the return and s_0 stands for the node at which the asset is traded. Consider an asset that pays $IV_{0\to1}(s_0)$ at the end of period 0. Empirically, such a variance claim can be constructed as a portfolio of options using the formula in Bakshi, Kapadia, and Madan [3].

At node 0_E, the variance of a one-period forward-looking return is $IV_{0\to1}(0_E) = Var[\ln P_1(s_1) - \ln P_0(0_E) | 0_E] = \phi^2\sigma^2$. At note 0_L, because investors anticipate no stochastic movement of the stock market in period 1, $IV_{0\to1}(0_L) = Var[\ln P_1(s_1) - \ln P_0(0_L) | 0_L] = 0$.

Then in period -1, the market value of this payoff can be computed as follows:

$$E\left[SDF(s_{-1}, s_0) \cdot IV_{0\to1}(s_0)\right] = \frac{1}{2}\beta\left(\frac{\bar{c}_0}{\bar{c}_{-1}}\right)^{-1}\frac{e^{-\frac{1}{\theta}V_0(0_E)}\phi^2\sigma^2}{E\left[e^{-\frac{1}{\theta}V_0(s_0)}\right]}, \tag{15}$$

where we assume that early and late resolution of uncertainty both happen with probability $\frac{1}{2}$. Clearly, the payoff of the above variance claim is increasing in informativeness. At node 0_E, investors expect that the stock market will respond to the upcoming announcement in period 1, and the option-implied variance is high. At node 0_L investors do not expect any stock market response to the (uninformative) announcement in the next period, and the option-implied variance is zero. The variance claim therefore satisfies the condition in Theorem 2 and can serve as a test asset.

To verify that the risk premium for the above variance claim identifies PER, denote the return an investor receives by holding this variance claim from period -1 to period 0 as $R_{Var}(s_{-1}, s_0)$. That is, $R_{Var}(s_{-1}, s_0)$ is the value of $IV_{0\to1}(s_0)$ divided by the price computed in (15). The risk premium on the return of the variance claim can be computed as:

$$\frac{E\left[R_{Var}(s_{-1}, s_0)\right]}{R_f(-1)} - 1 = \frac{1}{2}\left(e^{(1-\beta)\frac{\beta^2}{2\theta^2}\sigma^2} - 1\right). \tag{16}$$

Clearly, the risk premium is positive if and only if $\beta < 1$.

The above example illustrates the basic idea of Theorem 2: the risk premium for the variance claim from period -1 to period 0 is positive if and only if the representative agent prefers early resolution of uncertainty. The end of period -1 to period 0 is the period of resolution of information quality. In period -1, investors do not know whether the upcoming announcement in period 1 will be informative. Therefore, they assign a probability of $\frac{1}{2}$ to an informative announcement and a probability of $\frac{1}{2}$ to a completely uninformative announcement. In period 0, at node 0_E, early resolution realizes, and the agent expects the upcoming announcement to be informative. At node 0_L, late resolution realizes, and the announcement in period 1 is expected to be uninformative. In the empirical analyses below, we map this example to the data to implement an asset-market-based test for PER.

4 Empirical evidence

4.1 Key elements for identifying PER

To operationalize our thought experiment and use financial market data to test PER, we use monetary policy announcements made by the FOMC as our primary example of announcements that reveal macroeconomic uncertainty. In order to test PER, we need to identify the event of resolution of information quality (ROIQ) in the data and assets with payoffs increasing in informativeness. Below we summarize the four key elements of our identification exercise, which serve as a guide for the empirical sections that ensue.

Variations in informativeness First, we establish that the informativeness of FOMC announcements changes over time. We interpret an FOMC announcement as the announcement in period 1 in the example in Section 3. The thought experiment discussed in the last section requires the informativeness of the announcement to be stochastic. More informative announcements correspond to the case of early resolution and less informative announcements to the case of late resolution.

To demonstrate the time-varying informativeness of FOMC announcements, we show that the implied volatility reduction varies substantially across FOMC announcements. Intuitively, when the announcement is informative, the forward-looking implied volatility of the S&P 500 index will drop during the announcement, as macroeconomic uncertainty is resolved. The variation in the implied volatility reduction across FOMC announcements is therefore evidence of time-varying informativeness. To confirm the above intuition, we show that i) the implied volatility of the S&P 500 index on average drops significantly over FOMC announcements, and ii) such reduction shows substantial variation.

Predictability of informativeness Second, we demonstrate that the heterogeneity in informativeness is perceived by the market and construct a market-price-based measure of expected informativeness. The thought experiment in Section 3 requires that the market must be able to distinguish early resolution (node 0_E in Figure 2) from late resolution (node 0_L) so that expected asset payoffs can respond to the expected informativeness of the upcoming announcement. We establish this empirically by showing that market data successfully predict the amount of implied volatility reduction in the upcoming announcement. Specifically, we use the ratio of short-term versus long-term implied volatility, or the inverse slope of the term structure of implied volatility, as the predictor of the implied volatility drop. Our inverse slope variable is defined as

$$Inv\_Slope = \frac{IV^9}{IV^{90}}, \tag{17}$$

where IV^9, representing short-term implied stock market volatility, is the CBOE implied volatility index with 9 days to maturity, and IV^{90} is the CBOE implied volatility index with 90 days to maturity and represents long-term implied volatility.

Implied volatility from option prices may be affected by changes in the volatility of economic fundamentals (such as the volatility of aggregate productivity shocks) or by the informativeness of macroeconomic news. Variation in the volatility of economic fundamentals presumably happens at a much lower frequency, and its impact should extend beyond the few days around FOMC announcements. Therefore, fundamental economic volatility is likely to affect both short-term and long-term volatility. The inverse slope constructed above allows us to control for the volatility of economic fundamentals. We show in the next section that Inv_Slope has significant predictive power for the implied volatility reduction on FOMC announcement days.

Resolution of information quality The third element of our exercise is the identification of the period of resolution of information quality (ROIQ)—the period during which the market uncovers the informativeness of the upcoming FOMC announcements. The key to our identification exercise is the calculation of the risk premium earned by the test asset during the period of ROIQ. So, we first need to identify this period.

Empirically, we take advantage of the news data from RavenPack Analytics to demonstrate that the period of ROIQ is roughly five weekdays before FOMC announcements. We construct a direct measure of market attention to the Fed from the Fed-related news counts in RavenPack. On these five days, we find a strong, positive relation between change in inverse slope, which reflects market expectations of the informativeness of the upcoming announcement, and the attention measure constructed from RavenPack. However, such positive correlation does not exist in other periods. This procedure identifies the five days before FOMC announcements as the period of ROIQ because in the period of ROIQ, investors regularly form expectations about the informativeness of the upcoming FOMC announcement, and higher expected informativeness feeds into both higher market attention and a higher inverse slope. Outside the period of ROIQ this correlation does not exist because there is no reliably detectable change in market expectations about the informativeness of the upcoming announcement.

Premium for claims to market volatility Having identified the period of ROIQ, our final step is to estimate the risk premium earned by assets whose payoff is increasing in the informativeness of the announcements. We use claims to market volatility with short maturities. In the context of Figure 2, consider a claim to a stock market return variance that expires at the end of period 1. At node 0_E, the case of early resolution, the upcoming announcement is expected to be informative, and the market is expected to react to the announcement. Therefore at node 0_E, the expected volatility of the market return over the announcement is high. In contrast, at node 0_L, the case of late resolution, the upcoming announcement is expected to be uninformative. In this case, there is no news in period 1, and the expected volatility of the announcement return at node 0_L will be low. In fact, it is zero in this example. Hence, as we have explicitly shown in Section 3.2, a claim to this period-1 market return variance is a suitable testing asset for the purpose of detecting PER.

Motivated by the above observation, in the data, we use synthetic variance claims with maturities after the announcements as the test asset. To construct the claim to the stock market return variance, we follow Bakshi, Kapadia, and Madan [3], who show that under no arbitrage, the second moment of log security returns under the risk-neutral measure can be constructed from option prices in the following way:

$$\frac{1}{T}E^{RN}[(\ln(S_T) - \ln(S_t))^2] = \frac{e^{rT}}{T}\left(\int_0^{S_t} \frac{2(1-\ln(\frac{K}{S_t}))}{K^2} Put[K]dK + \int_{S_t}^{\infty} \frac{2(1+\ln(\frac{K}{S_t}))}{K^2} Call[K]dK\right) \tag{18}$$

Here, $E^{RN}[\cdot]$ is the risk neutral expectation, S_t is the price of the underlying security at time t and S_T the price at time T where $t < T$, and $Put[K]$ and $Call[K]$ are the prices of a put and call option with the underlying security S, strike price K, and expiration T. The formula expresses the risk-neutral squared log returns as integrals of options across strike prices. Because the squared mean of returns is orders of magnitude smaller than the mean of squared returns, Equation (18) practically measures the risk-neutral price of the return variance.

Empirically, we use the weighted sum of options with different strikes to approximate the above integral and construct the claims to the aggregate stock market variance. We also construct the at-the-money straddles as a robustness check. While variance claims closely align with our theory, straddles are simpler instruments that heavily load on volatility.

This empirical construct is consistent with the example in Section 3.2. As in the example, a more informative announcement is associated with a higher realized variance upon the announcement. Anticipating such a higher realized variance, the price of the variance claim in Equation (18) will rise during the period of ROIQ. As a result, the price of this variance claim is increasing in the informativeness of the upcoming announcement, and the risk premium of such asset can be used to detect PER.

We empirically estimate the excess return of the above portfolios during the period of ROIQ. Our Theorem 2 implies that an extra positive (negative) average return during the period of ROIQ is indicative of investors' preference for early (late) resolution of uncertainty.

4.2 Resolution of information quality

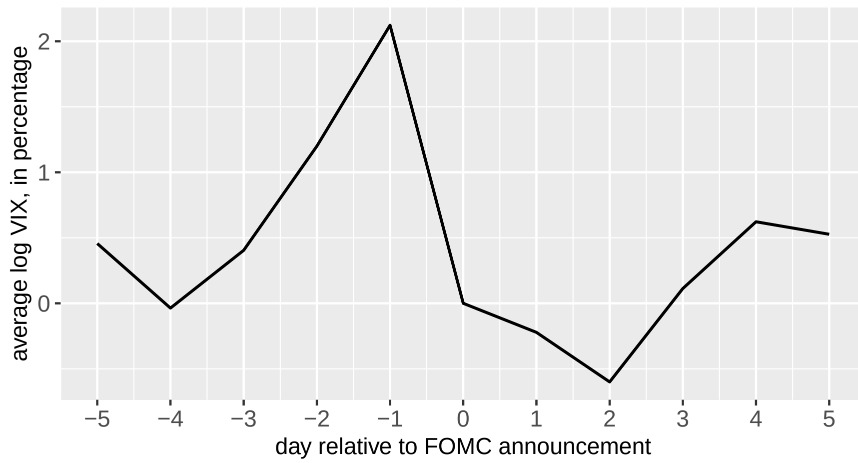
In this section, we first verify the four elements for the identification of PER we put forth in the last section. We then provide an estimation of the risk premium for the claim to the aggregate stock market variance in the period of ROIQ, which, according to Theorem 2, identifies investors' attitude toward the timing of resolution of uncertainty.

The option return data we use in our empirical exercises below come from OptionMetrics and are daily from 1996 to 2019. The implied volatility data we use include the 9-day, 30-day (VIX), and 90-day implied volatility indices on the S&P 500 from CBOE. The 30-day implied volatility is the VIX index, which goes back to 1990. The 9-day and 90-day IV indices have a shorter history going back to 2011 and 2007, respectively. These implied volatility indices end in 2020.

The 9-day implied volatility index has the shortest maturity. Therefore, the test asset in Equation (18) constructed using the 9-day implied volatility is less affected by measurement error induced by the volatility on non-announcement days. It, however, has a much shorter history than the 30-day implied volatility index. In what follows, we use the reduction in the 30-day implied volatility index as our baseline measure of realized informativeness and use the 9-day index as an alternative measure for robustness analysis.

Reductions in implied volatility across announcements In support of the hypothesis that FOMC announcements reduce uncertainty about the aggregate economy, we first show that on average there is a significant reduction in implied volatility on FOMC announcement days. The reduction in implied volatility is quite robust across all maturities. In Figure 5, we plot the level of the log VIX index around FOMC announcement days with the announcement-day log VIX normalized to zero. We denote the FOMC announcement day as day 0, the day before the announcement day as day -1, the day after as day 1, and so

Figure 5: **Log VIX around FOMC announcements**



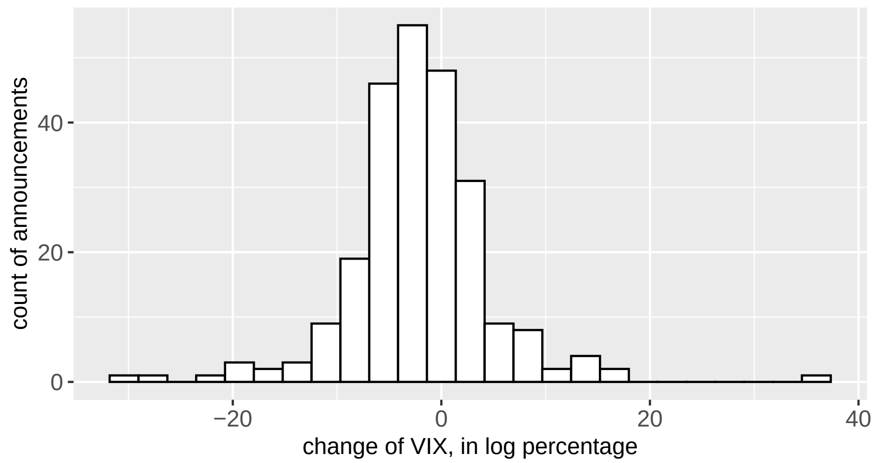
This figure illustrates the average log VIX index around FOMC announcements. We normalize the (end-of-day) log VIX index to zero for the FOMC announcement day, which is represented by day 0. Other days are labeled relative to the FOMC announcement day. The decline from 2.2 to 0 over day 0 means that the VIX index experiences on average a 2.2% decline on FOMC announcement days.

on. All values of the VIX index are end-of-the-day values. Figure 5 shows a clear reduction in VIX on FOMC announcement days on average. In Table 1, we present a formal regression analysis for the reduction in the VIX index on announcement days controlling for the day-of-the-week effect.[7] The third column is the reduction in the 30-day implied volatility and the fourth column is the reduction in the 9-day implied volatility. The reduction in the VIX index on announcement days is significant with a point estimate of -1.89%. Because the VIX index corresponds to the average volatility over 30 days, under the assumption that stock returns are i.i.d., a -1.89% reduction roughly corresponds to a 50% higher volatility on announcement days relative to non-announcement days.[8] The estimate for the 9-day implied volatility shows a similar pattern.

[7]As shown in Table 1, the VIX index has a significant day-of-the-week pattern. In particular, changes in the VIX index are typically positive on Mondays and negative on Wednesdays and Fridays. Because FOMC announcements are not evenly distributed across days of the week, we control for this effect out of an abundance of caution.

[8]Assume that the daily volatility is σ on non-announcement days and $(1+x)\,\sigma$ on announcement days. The 30-day volatility before announcements is $\sqrt{(1+x)^2\,\sigma^2+29\sigma^2}$, and the 30-day volatility after announcement is $\sqrt{30\sigma^2}$. A log difference of 2% between the above volatility measures translates into a value of $x = 49\%$.

Figure 6: **Histogram of changes in log VIX index on FOMC announcement days**



This figure plots the histogram of changes in the log VIX index around FOMC announcements. Changes in the log VIX index are computed as the difference between the log of the VIX index at the end of the announcement day and that on the day before the announcement day.

Our identification exercise requires that the informativeness of FOMC announcements be time varying. Here, we provide consistent empirical evidence by demonstrating that the amount of volatility reduction shows substantial variations across announcements. We plot the histogram for the changes in the VIX index on FOMC announcement days in Figure 6. Note the fairly wide range of implied volatility changes across announcements, indicating that the informativeness of announcements does change over time.

Predictability of informativeness The second element of our identification exercise is the predictability of informativeness. We establish this by demonstrating that the reduction in volatility across announcements can be predicted by the inverse slope of the term structure of implied volatility. To do this, we regress the changes in short-term implied volatility on the inverse slope of the previous day, an FOMC announcement day dummy, an interaction between the two terms, and control variables such as the day-of-the-week dummies:

$$\Delta \ln IV_t = \xi_0 + \xi_1 Inv\_Slope_{t-1} + \xi_2 I_t^{FOMC} + \xi_3 Inv\_Slope_{t-1} \cdot I_t^{FOMC} + \sum_{d=1}^{5} \delta_d I_{d,t}^{DOW} + \varepsilon_t. \quad (19)$$

Here, $\Delta \ln IV_t$ is the one-day change in implied volatility from the end of $t-1$ to t, Inv_Slope_{t-1} is the inverse slope defined in Equation (17) on day $t-1$, and I_t^{FOMC} is an indicator variable that takes the value of 1 if day t is a pre-scheduled FOMC announcement day. For $d = 1, 2, \cdots, 5$, $I_{d,t}^{DOW}$ is an indicator variable that takes the value of 1 if day t is the dth day of the week. As explained earlier, we expect short-term volatility to be higher relative to long-term volatility ahead of informative FOMC announcements because higher informativeness of announcements, if anticipated by the market, should be associated with larger reactions of stock market returns with respect to these announcements.

In Table 2, we report several versions of the above regression to demonstrate the predictability of announcement-day volatility reductions. In column (1), the regression of the volatility reduction on the inverse slope produces a significant coefficient of -6.57, indicating that, in general, the inverse slope variable has significant predictive power for volatility reductions. It is well known that volatility is mean reverting. As shown in column (2), higher volatility on the previous day is associated with significantly larger volatility reductions as well. However, whenever the inverse slope variable is included (columns (3), (4) and (5)), the effect of the level of volatility on the previous day is subsumed. The regression in column (4) includes only the 77 observations on FOMC announcement days. In this case, the effect of the inverse slope is much larger in magnitude, although the t-statistic is much smaller because of a much smaller sample. In column (5), we report the result of the full regression. Here, Inv_Slope_{t-1} has significant predictive power for implied volatility reductions in general. More importantly, the coefficient on the interaction term of the FOMC indicator and Inv_Slope_{t-1} is significantly larger, indicating that the Inv_Slope_{t-1} variable has extra predictive power on FOMC announcement days. In the last column of the same table, we report the results of regression (19), where the dependent variable is the reduction in the 9-day implied volatility. This regression shows a similar pattern with a more negative point estimate for ξ_3. These results indicate that the option market correctly understands the informativeness of the FOMC announcements ahead of time and expresses its view via option prices. Anticipating an informative announcement, investors bid up the prices of the short-horizon options relative to long-horizon ones, creating a large 9-day/90-day implied volatility ratio before the announcement. In the next section, we investigate over which periods investors come up with this expected informativeness.

Period of resolution of information quality The third step of our identification exercise is to identify the period of resolution of information quality. As explained earlier, we do this by first constructing a time series that measures the market's attention to the Fed.

We obtain the number of Fed-related new items from RavenPack Analytics. The measure is the number of these news items issued on a given day divided by the average number in the past 30 days. This division step keeps the measure stationary, whereas the number of news items has an upward trend over time. We refer to this ratio as the news intensity.

Column (1) of Table 3 performs a daily time-series regression of the news intensity measure on the contemporaneous daily change in the inverse slope. It shows that on average, Fed-related news intensity does not strongly correlate with the inverse slope measure. As explained earlier, there is no reason to expect these two measures to positively correlate, except during the period of ROIQ. Column (2)-(6) perform the same regression on various subsamples around FOMC announcement days. Column (2)-(3) show that during the five weekdays before the FOMC announcements, the two measures are significantly positively correlated. This finding suggests that during these five days, investors regularly form expectations of the informativeness of the upcoming FOMC announcement (note that this is the definition of the period of ROIQ), and higher expected informativeness corresponds to both a higher inverse slope and more news about the Fed. Columns (4)-(6) are placebo tests showing that there is no positive relation between the two time series on and after FOMC announcement days. The negative coefficient on FOMC announcement days is because more informative announcements see a higher reduction in the inverse slope and also receive more news attention.

4.3 The PER premium

The last step of our identification exercise is the estimation of the premium of the claim to market volatility constructed in Equation (18). Theorem 2 implies that if investors have preference for early resolution, this premium must be positive during the period of ROIQ. In the data, we must also take into account the premium that variance claims normally receive—both within and outside the period of ROIQ. On an average day, variance claims provide a valuable hedge against stock market crashes and adverse economic shocks in general. It is therefore unsurprising that they receive a negative premium on average. Such protection exists both within and outside the period of ROIQ. Consequently, we should not seek an outright positive premium on variance claims over the period of ROIQ but rather a positive premium relative to an average day.

To estimate the sign of this PER premium, we construct synthetic variance claims on the S&P 500 index using put and call prices from OptionMetrics, the range of which is 1996 to 2019. The claims are constructed according to Equation (18) and are portfolios of out-of-money puts and calls. The construction details can be found in the data appendix. We also construct the at-the-money straddles. With daily returns to these variability-paying portfolios, we run the following regression:

$$r_{\tau,t} = \beta I_t^{ROIQ} \cdot I_t^{After}(\tau) + \beta_1 I_t^{FOMC} + \sum_{w=1}^{11} \gamma_w I_{w,t}^{Maturity}(\tau) + \sum_{d=1}^{5} \delta_d I_{d,t}^{DOW} + \epsilon_{\tau,t}. \tag{20}$$

This is a panel regression where $r_{\tau,t}$ is the log return realized on date t on a claim to market volatility constructed using an options portfolio with maturity τ, I_t^{ROIQ} is an indicator function that takes the value of 1 if date t is within the period of ROIQ of a pre-scheduled FOMC announcement, and $I_t^{After}(\tau)$ is an indicator function that takes the value of 1 if the claim expires after the closest announcement in the future as of day t. Because the price of options that expire before announcements will not be affected by the informativeness of these announcements, we focus only on options that expire after the announcements. Here, I_t^{FOMC} is an indicator function that takes the value of 1 if day t is an FOMC announcement day. We also include several control variables in the above regression: $I_w^{Maturity}(\tau)$ is an indicator function for the maturity of the options, which takes the value of 1 if the option is w weeks to maturity, for $w = 1, 2, \cdots, 11$. As before, $I_{d,t}^{DOW}$ are indicator variables that control for the day-of-the-week effect.

We present our regression results in Table 4, where we report the coefficient β, which captures the average return of the variance claims over the period of ROIQ in excess of their returns on an average day. Column (1) reports the excess return on the second moment portfolios and column (2) that on the at-the-money straddles.

What is important for our theory is the coefficient β, which is what Table 4 shows. In both columns, we observe a significantly positive coefficient. This finding indicates that these variance-paying portfolios see high excess returns over the period of ROIQ, consistent with a preference for early resolution of uncertainty.

It is worth mentioning that, first, these portfolios do not have a higher loading on market excess returns over the period of ROIQ. Table 5 shows, if anything, that the market loading is somewhat lower. Second, the market return is not higher during the period of ROIQ. In fact, over this period, the market return is about 8 basis points lower than average. Given these two empirical patterns, this premium on the variability-paying portfolios cannot be driven by exposure to the market. Controlling for the market or the Fama-French three factors in the regression of Table 4 does not appreciably change the coefficient β. This robustness check is useful because an important assumption that we make in our analysis is that the period of ROIQ reveals the informativeness of the upcoming announcement, but not the news in the announcement. The assumption seems consistent with the data because i) the market itself does not earn a positive premium over the period of ROIQ, and ii) the premium of the variance claims over the period of ROIQ is not explained by the market or common risk factors.

4.4 Additional Results

A simple, commonly used instrument of variance claims is the VIX futures, which pay the level of the VIX index on the expiration day. While the history of these futures is relatively short, they are simple instruments that load on volatility. However, notice that because they pay the VIX level as of the expiration day, a VIX future that expires after an FOMC announcement is not exposed to the volatility over the announcement, which captures the informativeness that is at the core of our theory. This is because the VIX index is a forward-looking index that captures the expected volatility over the 30 days in the future and not the past. The VIX futures therefore enable us to conduct a valuable placebo test.[9]

Table 6 repeats the regression from Table 4, except that now the dependent variables are log returns on VIX futures. The table shows that VIX futures have no significant excess returns during the period of ROIQ relative to an average day. This test is valuable because VIX futures are similar to our synthetic variance claims in nature, but the subtle difference of being forward-looking predicts that our theory should not apply to them. This evidence lends further support to our theory by showing that the ROIQ premium pattern we see on variance claims really stems from the exposure to market movements during the announcements.

While the FOMC announcements clearly resolve important systematic risks, there are relatively few observations along these lines. Savor and Wilson (2016) demonstrate that individual firms' earnings announcements also resolve important systematic cash flow risks. Within the 3-day window centering on the earnings announcement day, a stock earns on average 25.8 basis points in excess of the market.[10] The economic scale of this risk premium is smaller than that earned by the aggregate market on FOMC announcement days but is on the same order of magnitude. Furthermore, investors and analysts pay close attention to these earnings announcements and make forecasts about the earnings outcome ahead of

[9]We thank Ian Dew-Becker for suggesting such a test in his discussion of this paper.

[10]The 25.8 basis point mean is weighted by the market value of the stock divided by the total market values of all stocks in the cross section of the CRSP universe from 1971 to 2021.

time.[11] Additionally, firms' management exercises considerable discretion in being vague or precise during earnings calls, much like the Fed does during FOMC announcements. A period of ROIQ may therefore also exist for these individual earnings announcements. Since a 9-day implied volatility index for individual stock options cannot be constructed, we cannot perform an analogous search for the period of ROIQ in this context. We therefore keep using 5 weekdays before the announcement and investigate whether the returns to the variance-paying portfolios—now on individual stock options—are also abnormally high before the earnings announcements. Table 7 shows exactly this.[12] This piece of evidence lends additional support to our results.

5 Conclusion

This paper develops a revealed preference theory for preference for the timing of resolution of uncertainty based on asset pricing data and presents corresponding empirical evidence. Our main theorem provides an equivalent characterization of the representative agent's preference for early resolution of uncertainty in terms of the risk premium of assets realized during the period of resolution of information quality of macroeconomic announcements. Empirically, we find support for preference for early resolution of uncertainty based on evidence on the dynamics of the implied volatility of S&P 500 index options before FOMC announcements.

[11] A systematic dataset containing these forecasts is the I/B/E/S database, available at Wharton Research Data Services (WRDS).

[12] Johnson and So (2018) show that the cost of trading negative news on stocks increases before earnings announcements and that this leads to an increase in stock prices prior to announcements. This would lead to elevated call prices and decreased put prices prior to the announcements because they embed long and short positions in stocks, respectively. However, because our variance-paying portfolios roughly equally weight puts and calls, these effects should largely cancel out each other.

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Table 1

Changes in VIX index on FOMC announcement days

	(1)	(2)	(3)	(4)
	$\Delta \ln VIX$	$\Delta \ln VIX$	$\Delta \ln VIX$	$\Delta \ln VIX9$
I_1^{DOW}	1.94***		1.94***	5.99***
	[10.20]		[10.20]	[8.86]
I_2^{DOW}	-0.26		-0.13	0.32
	[-1.59]		[-0.77]	[0.64]
I_3^{DOW}	-0.48***		-0.33**	-1.06*
	[-3.17]		[-2.12]	[-1.86]
I_4^{DOW}	-0.04		-0.03	-0.56
	[-0.22]		[-0.17]	[-1.05]
I_5^{DOW}	-1.00***		-1.00***	-3.74***
	[-5.91]		[-5.90]	[-7.01]
I_{FOMC}		-2.20***	-1.89***	-2.43*
		[-5.02]	[-4.18]	[-1.71]
N	7,766	7,766	7,766	2,477
R^2	0.022	0.003	0.024	0.063

This table reports results from running the following daily time-series regression: $\Delta \ln IV_t = \sum_{d=1}^{5} \delta_d I_{d,t}^{DOW} + \xi I_t^{FOMC} + \epsilon_t$, where $\Delta \ln IV_t$ is the change in $\ln VIX$ on day t (in percentage units), $I_{d,t}^{DOW}$ is the indicator of whether day t is the dth weekday (e.g., $I_{1,t}^{DOW}$ takes the value of 1 when day t is Monday, and 0 otherwise), and I_t^{FOMC} is the indicator of whether day t is an FOMC announcement day. The dependent variable in columns (1)-(3) is based on the 30-day VIX, and that in column (4) is based on the 9-day VIX. Data are daily from 1990 to 2020 in columns (1)-(3) and from 2011 to 2020 in column (4). T-statistics are computed with White standard errors and reported in square brackets.

Table 2

Predictability of implied volatility reduction on FOMC announcement days

	(1)	(2)	(3)	(4)	(5)	(6)
	$\Delta \ln VIX$	$\Delta \ln VIX$	$\Delta \ln VIX$	$\Delta \ln VIX$	$\Delta \ln VIX$	$\Delta \ln VIX9$
Inv_Slope_{t-1}	-6.57***		-5.37***	-15.84*	-6.13***	-14.47***
	[-4.23]		[-2.85]	[-1.87]	[-3.94]	[-6.65]
VIX_{t-1}		-0.11***	-0.04	-0.13		
		[-4.11]	[-1.24]	[-0.64]		
I_t^{FOMC}					10.87	16.57**
					[1.58]	[2.08]
$Inv\_Slope_{t-1} \cdot I_t^{FOMC}$					-13.42*	-19.90**
					[-1.73]	[-2.21]
DOW Indicators	Yes	Yes	Yes	No	Yes	Yes
Constant	No	No	No	Yes	No	No
N	2477	2477	2477	77	2477	2477
R^2	0.035	0.029	0.036	0.139	0.038	0.101

Column (5) of this table reports results from running the following daily time-series regression: $\Delta \ln IV_t = \xi_0 + \xi_1 Inv\_Slope_{t-1} + \xi_2 I_t^{FOMC} + \xi_3 Inv\_Slope_{t-1} \cdot I_t^{FOMC} + \sum_{d=1}^{5} \delta_d I_{d,t}^{DOW} + \varepsilon_t$, where $\Delta \ln IV_t$ is the change in log VIX on day t (in percentage units), Inv_Slope_{t-1} is the inverse slope, or the 9-day VIX divided by the 30-day VIX, on day $t-1$, I_t^{FOMC} is the indicator of whether day t is an FOMC announcement day, and $I_{d,t}^{DOW}$ are indicators of whether day t is the dth weekday (e.g., $I_{1,t}^{DOW}$ takes the value of 1 when day t is Monday, and 0 otherwise). Columns (1) to (3) are the regressions with subsets of the independent variables and possibly adding VIX_{t-1}, which is the VIX level of day $t-1$. Column (4) is restricted to FOMC announcement days only. Column (6) has a different dependent variable, which is the change in the 9-day VIX. Data are daily from 2011 to 2020. T-statistics are computed with White standard errors and reported in square brackets.

Table 3

News intensity and change in inverse slope over the period of ROIQ

	$News\_Intensity_t$					
	All t	$[-10,-6]$	[-5,-1]	FOMC day	[1,5]	[6,10]
ΔInv_Slope_t	-0.242	-0.209	1.076***	-2.132**	0.098	-0.382
	[-1.13]	[-0.69]	[4.09]	[-2.02]	[0.16]	[-1.50]
N	2,453	385	385	77	385	385

Column 1 of this table reports the results of the following time-series regression: $News\_Intensity_t = \alpha + \beta\Delta Inv\_Slope_t + \epsilon_t$. Here, $News\_Intensity_t$ is the number of Fed-related news items on day t divided by the average number of items in the past 30 days. Inv_Slope_t is the 9-day VIX divided the by 90-day VIX. Columns (2) and (3) perform the same regression conditioning on day t being 10 to 6 and 5 to 1 weekdays, inclusive, before the FOMC announcements. Column (4) is on the FOMC announcement days. Columns (5) to (6) are 5 weekdays and 6 to 10 weekdays after the FOMC announcements. Data are daily from 2011 to 2020. T-statistics are computed using White standard errors and reported in square brackets.

Table 4

Excess returns of stock index options during the period of ROIQ

	(1)	(2)
	2nd Moment	Straddle
ROIQ premium β	1.085**	0.428***
	[2.44]	[2.25]
N	41,982	41,991
R^2	0.101	0.054

This table reports the results of the following panel regression: $r_{\tau,t} = \beta I_t^{ROIQ} \cdot I_t^{After}(\tau) + \beta_1 I_t^{FOMC} + \sum_{w=1}^{11} \gamma_w I_{w,t}^{Maturity}(\tau) + \sum_{d=1}^{5} \delta_d I_{d,t}^{DOW} + \epsilon_{\tau,t}$, where $r_{\tau,t}$ is the log return of the option portfolio with expiration τ on day t (in percentage units), I_t^{ROIQ} is an indicator of whether day t is within the period of ROIQ, $I_t^{After}(\tau)$ is an indicator of whether τ is after the next FOMC announcement as of day t, I_t^{FOMC} indicates whether day t is an FOMC announcement day, $I_{w,t}^{Maturity}(\tau)$ is an indicator of whether t is within w weeks of τ, and $I_{d,t}^{DOW}$ is an indicator of whether day t is the dth weekday (e.g., $I_{1,t}^{DOW}$ takes the value of 1 when day t is Monday, and 0 otherwise). Regressions apply equal weight on each trading day (i.e., t). Column (1) includes portfolios tracking the second moment of the underlying returns, and column (2) includes the at-the-money straddles. The at-the-money strike price is the one closest to the underlying index level. Data are daily from 1996 to 2019. T-statistics are computed using clustered standard errors by trading day and reported in square brackets.

Table 5

Market loadings of options prior to FOMC announcements

	(1)	(2)
	2nd Moment	Straddles
β_{mkt}	-7.492***	-1.122***
	[-24.18]	[-6.09]
β_{mkt}^{ROIQ}	0.591	0.217
	[0.72]	[0.50]
N	41,982	41,991
R^2	0.246	0.054

This table reports the results of the following panel regression: $r_{\tau,t} = \beta_{mkt} mkt_t + \beta_{mkt}^{ROIQ} mkt_t \cdot I_t^{ROIQ} \cdot I_t^{After}(\tau) + \beta I_t^{ROIQ} \cdot I_t^{After}(\tau) + \beta_1 I_t^{FOMC} + \sum_{w=1}^{11} \gamma_w I_{w,t}^{Maturity}(\tau) + \sum_{d=1}^{5} \delta_d I_{d,t}^{DOW} + \epsilon_{\tau,t}$, where $r_{\tau,t}$ is the log return of the option portfolio with expiration τ on day t (in percentage units), mkt_t is the log market return on day t in excess of the risk-free rate, I_t^{ROIQ} is an indicator of whether day t is within the period of ROIQ, $I_t^{After}(\tau)$ is an indicator of whether τ is after the next FOMC announcement as of day t, I_t^{FOMC} is an indicator of whether day t is an FOMC announcement day, $I_{w,t}^{Maturity}(\tau)$ is an indicator of whether t is within w weeks of τ, and $I_{d,t}^{DOW}$ is an indicator of whether day t is the dth weekday (e.g., $I_{1,t}^{DOW}$ takes the value of 1 when day t is Monday, and 0 otherwise). Regressions apply equal weight on each trading day. Column (1) uses portfolios tracking the second moment of the underlying returns, and column (2) uses at-the-money straddles. Data are daily from 1996 to 2019. T-statistics are computed using clustered standard errors by trading day and reported in square brackets.

Table 6

Excess returns of VIX futures during the period of ROIQ

	VIX futures
ROIQ premium β	0.033
	[0.24]
N	10,598
R^2	0.007

This table reports the results of the following panel regression: $r_{\tau,t} = \beta I_t^{ROIQ} \cdot I_t^{After}(\tau) + \beta_1 I_t^{FOMC} + \sum_{w=1}^{11} \gamma_w I_{w,t}^{Maturity}(\tau) + \sum_{d=1}^{5} \delta_d I_{d,t}^{DOW} + \epsilon_{\tau,t}$, where $r_{\tau,t}$ is the log return of the VIX future with expiration τ on day t (in percentage units), I_t^{ROIQ} is an indicator of whether day t is within the period of ROIQ, $I_t^{After}(\tau)$ is an indicator of whether τ is after the next FOMC announcement as of day t, I_t^{FOMC} indicates whether day t is an FOMC announcement day, $I_{w,t}^{Maturity}(\tau)$ is an indicator of whether t is within w weeks of τ, and $I_{d,t}^{DOW}$ is an indicator of whether day t is the dth weekday (e.g., $I_{1,t}^{DOW}$ takes the value of 1 when day t is Monday, and 0 otherwise). Regressions apply equal weight on each trading day (i.e., t). Data are daily from 2004 to 2019. T-statistics are computed using clustered standard errors by trading day and reported in square brackets.

Table 7

Excess returns of stock options prior to earnings announcements

	(1)	(2)
	2nd Moment	Straddle
ROIQ premium β	1.276***	0.773***
	[6.62]	[9.86]
N	660,519	1,301,995
R^2	0.028	0.013

This table reports the results of the following panel regression: $r_{\tau,i,t} = \beta I_{i,t}^{ROIQ} \cdot I_{i,t}^{After}(\tau) + \beta_1 I_{i,t}^{EA} + \sum_{w=1}^{11} \gamma_w I_{w,t}^{Maturity}(\tau) + \sum_{d=1}^{5} \delta_d I_{d,t}^{DOW} + \epsilon_{\tau,t}$, where $r_{\tau,i,t}$ is the log return of the option portfolio of stock i with expiration τ on day t (in percentage units), $I_{i,t}^{ROIQ}$ is an indicator of whether day t is within the period of ROIQ for stock i, $I_{i,t}^{After}(\tau)$ is an indicator of whether τ is after the next earnings announcement as of day t, $I_{i,t}^{EA}$ indicates whether day t is an earnings announcement day for stock i, $I_{w,t}^{Maturity}(\tau)$ is an indicator of whether t is within w weeks of τ, and $I_{d,t}^{DOW}$ is an indicator of whether day t is the dth weekday (e.g., $I_{1,t}^{DOW}$ takes the value of 1 when day t is Monday, and 0 otherwise). Regressions apply equal weight on each trading day (i.e. t), and restrict to the S&P 500 universe. Column (1) includes portfolios tracking the second moment of the underlying returns, and (2) are includes at-the-money straddles. The at-the-money strike price is the one closest to the underlying index level. Column (1) requires that there are at least 10 instruments in the portfolio and column (2) requires that the at-the-money strike price is chosen from at least 10 different strike prices, and is neither the maximum nor the minimum among them. Data are daily from 1996 to 2019. T-statistics are computed using clustered standard errors by trading day and reported in square brackets.

Appendix (For online publication)

A Proof for Theorems 1 and 2

Proof for Theorem 1 Because the underlying probability space Ω is finite dimensional, for any random variable V defined on Ω, we can identify V as a finite dimensional vector $V = [V(1),\ V(2),\ \cdots,\ V(n)]$ and think of the certainty equivalent functional $\mathcal{I}$ as a function from $\boldsymbol{R}^n$ to $\boldsymbol{R}$. For $s = 1,\ 2,\ \cdots,\ n$, we denote $\frac{\partial}{\partial V(s)}\mathcal{I}[V]$ as the partial derivative of $\mathcal{I}$ with respect to the sth element of V. The stochastic discount factor can be computed from the marginal rate of substitution of the representative agent. Given the form of the utility function in (1), the SDF is given by:

$$SDF(s_0, s_1) = \beta \frac{1}{\mu(s_1)} \frac{\frac{\partial \mathcal{I}[V_1]}{\partial V_1(s_1)} u'(\bar{c}_1)}{u'(\bar{c}_0)} = \lambda \frac{\partial \mathcal{I}[V_1]}{\partial V_1(s_1)}, \tag{21}$$

where $\lambda = \beta \frac{1}{\mu(s_1)} \frac{u'(\bar{c}_1)}{u'(\bar{c}_0)}$ is a constant that does not depend on s_1. Recall that $\mu(s_1) = \frac{1}{n}$ for all s_1 due to the assumption of equal probability.

To prove Theorem 1, we first set up some notation and introduce a useful lemma. Note that given the SDF, no arbitrage implies that the price of any period-1 payoff X denominated in period-0 consumption goods is given by $P_0(X) = E_0[SDF(s_0, s_1)X(s_1)]$. The one-period risk-free rate paid in period 1 is $R_f(0) = \frac{1}{E_0[SDF(s_0,s_1))]}$. The risk-premium for an asset with payoff X is therefore given by $E_0\left[\frac{X}{P_0(X)}\right] - R_f(0)$.

Lemma 1. *Suppose that $\mathcal{I} : \mathcal{L}(\Omega, \mathcal{F}, P) \to \boldsymbol{R}$ is strictly increasing and continuously differentiable. The following conditions are equivalent:*

(i) The risk premium received in period 1 is non-negative for all payoffs that are comonotone with respective to V_1.

(ii) $\mathcal{I}$ is non-decreasing in second order stochastic dominance, that is, $\forall\ V$ and $\tilde{V} \in \mathcal{L}(\Omega, \mathcal{F}, P)$, if V second order stochastic dominates $\tilde{V}$ then $\mathcal{I}[V] \geq \mathcal{I}\left[\tilde{V}\right]$.

(iii) For any $V \in \mathcal{L}(\Omega, \mathcal{F}, P)$,

$$\left[\frac{\partial}{\partial V(s)}\mathcal{I}[V] - \frac{\partial}{\partial V(s')}\mathcal{I}[V]\right][V(s) - V(s')] \leq 0. \tag{22}$$

Proof. Here, we prove the equivalence between statements (i) and (iii). The equivalence between (ii) and (iii) is based a characterization of Schur concavity that can be found in Marshall, Arnold, and Olkin [37] or Muller and Stoyan [39].

First, we assume that statement (i) is true and prove (iii) by contradiction. Suppose there exists $V \in \mathcal{L}(\Omega, \mathcal{F}, P)$ and s, s' such that

$$V(s) > V(s'), \; and \; \frac{\partial}{\partial V(s)} \mathcal{I}[V] > \frac{\partial}{\partial V(s')} \mathcal{I}[V]. \tag{23}$$

Consider the following payoff:

$$X(i) = V(i) \quad for \quad i = s_1, s_1'; \quad X(i) = 0 \quad otherwise.$$

Given condition (23), X is strictly positively correlated $\frac{\partial \mathcal{I}[V_1]}{\partial V_1(s_1)}$ and, therefore, the SDF defined in (21). As a result,

$$P_0(X) = E[SDF(s_0, s_1) X(s_1)] > E[SDF(s_0, s_1)] E[X(s_1)] = \frac{E[X(s_1)]}{R_f(0)},$$

That is, the risk premium for X is strictly negative. However, by the definition of comonotonicity in equation (12), X is comonotone with V_1, a contradiction.

Next, we assume that statement (iii) in the lemma is true and prove (i). Take any X that is comonotone with V_1. By condition (22), X is negatively comonotone with respect to $\frac{\partial \mathcal{I}[V_1]}{\partial V_1(s_1)}$ and the SDF defined in (21). As a result, X and SDF are negatively correlated and

$$P_0(X) = E[SDF(s_0, s_1) X(s_1)] \leq E[SDF(s_0, s_1)] E[X(s_1)] = \frac{E[X(s_1)]}{R_f(0)}$$

as needed. $\square$

It is straightforward to show that the strict inequality version of Lemma 1 also holds. That is, the following under the same assumptions in Lemma 1, the following states are also equivalent:

(i') The risk premium received in period 1 is strictly positive for all payoffs that are strictly comonotone with respective to V_1.

(ii') $\mathcal{I}$ is strictly increasing in second order stochastic dominance, that is, $\forall$ V and $\tilde{V} \in \mathcal{L}(\Omega, \mathcal{F}, P)$, if V strictly second order stochastic dominates $\tilde{V}$ then $\mathcal{I}[V] > \mathcal{I}\left[\tilde{V}\right]$.

(iii') For any $V \in \mathcal{L}(\Omega, \mathcal{F}, P)$,

$$\left[\frac{\partial}{\partial V(s)}\mathcal{I}[V] - \frac{\partial}{\partial V(s')}\mathcal{I}[V]\right][V(s) - V(s')] \leq 0. \tag{24}$$

and the strict inequality holds as long as $V(s) \neq V(s')$.

To prove Theorem 1, we note that statement 2 in Theorem 1 is equivalent to statement (ii) in Lemma 1. In addition, statement 3 in Theorem 1 is equivalent to statement (iii) in Lemma 1. It is enough to show that statement 1 is equivalent to (i). Given that u is a strictly increasing function, the definition of V_1 in equation (11) implies that $c_2(s_1)$ is strictly comonotone with $V_1(s_1)$. This establishes the equivalence between statement 1 in Theorem 1 and statement (i) in Lemma 1. The strict inequality version of the theorem can be similarly proved by using the strict inequality version of Lemma 1.

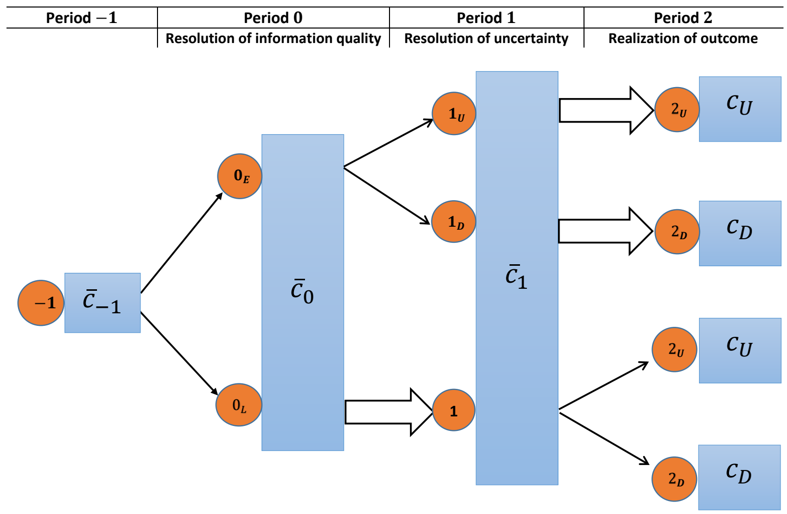
Proof for Theorem 2 First, we assume condition 1 in Theorem 1 is true, that is the risk premium for any asset with payoff comonotone with informativeness is non-negative. To prove condition 2, it is enough to show that $V_0(s_0)$ is comonotone with informativeness. We prove by contradiction. Assume $\exists s_0$ and s_0' such that $\iota(s_0) < \iota(s_0')$ and $V_0(s_0) < V_0(s_0')$. Consider the following payoff:

$$X(i) = \frac{1}{\iota(i)} \quad if \quad i = s_0, s_0'; \; X(i) = 0 \; otherwise. \tag{25}$$

Clearly, X is comonotone with informativeness. By condition 1, the risk premium of X must be non-negative. Note that X is also strictly negatively comonotone with $V_0(s_0)$. By Lemma 1, we know that under the assumption of strict GRS, the risk premium for X must be strictly negative, which is a contradiction.

Next, we assume that condition 2 in Theorem 1 holds and prove condition 1. Note that preference for early resolution of uncertainty is equivalent to $V_0(s_0)$ being comonotone with respect to informativeness. As a result, any payoff that is comonotone with respect to informativeness is also comonotone with $V_0(s_0)$. By the assumption of GRS, we know that the risk premium on this asset must be non-negative. The strict inequality version of this theorem can be proved similarly.

B Example



B.1 Notations

We adopt the following general convention for notation. We use subscripts for the time period, and write in parenthesis the node at which the quantity or price is calculated. Note that all nodes in period -1, 0, and 1 in Figure 2, reproduced above, has a unique name. Below we first set up the notations.

Utility

- Period -1: V_{-1} is the utility in period -1.
- Period 0: There are two nodes in period 0. The corresponding utilities are $V_0(0_E)$ and $V_0(0_L)$, or $V_0(s_0)$ in general.
- Period 1: There are three nodes in period 1. The corresponding utilities are $V_1(1_U)$, $V_1(1_D)$, and $V_1(1)$, or $V_1(s_1)$ in general.

SDF We also compute the SDF that prices one-period cash flows. We use the following convention: $SDF(node_t, node_{t+1})$, which prices period $t+1$ cash flow into period t consumption units (note that we are looking at one-period SDFs). To be precise, the notation $(node_t, node_{t+1})$ emphasizes that the SDF computes the price of a cash flow delivered at

$node_{t+1}$ in terms of $node_t$ consumption units. In this notation, $node_{t+1}$ must be a node that immediately follows $node_t$.

- Pricing period 0 cash flow into period -1 consumption units: $SDF(-1, s_0)$. There is only one node in period -1, and $SDF(-1, 0_E)$, $SDF(-1, 0_L)$ are the realizations of the random variable $SDF(-1, s_0)$.

- Pricing period 1 cash flow into period 0 consumption units: $SDF(0_L, s_1)$ and $SDF(0_E, s_1)$. There are two nodes in period 0. For 0_E, $SDF(0_E, 1_U)$ and $SDF(0_E, 1_D)$ are the two realizations of the random variable $SDF(0_E, s_1)$. At node 0_L, $SDF(0_L, 1)$ is the only possible value, as there is only one following node.

- Pricing period 2 cash flow into period 1 consumption units: $SDF(1_U, s_2)$, $SDF(1_D, s_2)$, and $SDF(1, s_2)$. There are three nodes in period 1. At node 1_U: $SDF(1_U, 2_U)$ is the only possible realization of the SDF, as there is only one following node. At node 1_D it is $SDF(1_D, 2_D)$, as again there is only one following node. At node 1: $SDF(1, 2_U)$ and $SDF(1, 2_D)$ are the two realizations of the random variable $SDF(1, s_2)$.

Price Here, we calculate the price of a levered consumption claim paid in period 2: $C_2^{1+\phi}$. The prices below refers to the price of this consumption claim evaluated at different nodes. We continue to adopt the same notation convention for evaluating utilities. That is, we will use subscripts for the time period, and we will write in parenthesis the node at which the prices are calculated.

- Period -1: P_{-1}.

- Period 0: $P_0(0_E)$, $P_0(0_L)$.

- Period 1: $P_1(1_U)$, $P_1(1_D)$, $P_1(1)$.

Variance claims We calculate the price of a claim to stock market returns variance. The stock market refers to the claim to the levered consumption as described above. We use subscripts for the horizon of the variance, and we write in parenthesis the node at which the expected variance is calculated. In general, $IV_{t\to t+1}(s_t)$ denotes the s_t expectation of the return variance from the end of period t to the end of $t+1$.

B.2 Utility calculation

In period 1:

$$\begin{aligned} V_1(1_U) &= \ln \overline{c_1} + \beta \ln c_2 \\ V_1(1_D) &= \ln \overline{c_1} - \beta\theta \ln E[e^{-\frac{1}{\theta}\ln c_2}] \\ &= \ln \overline{c_1} + \beta(\mu - \frac{1}{2}\frac{1}{\theta}\sigma^2) \end{aligned}$$

In period 0:

$$\begin{aligned} V_0(0_E) &= \ln \overline{c_0} - \beta\theta \ln E[e^{-\frac{1}{\theta}V_1(s_1)}] \\ &= \ln \overline{c_0} - \beta\theta \ln E[e^{-\frac{1}{\theta}(\ln \overline{c_1} + \beta \ln c_2)}] \\ &= \ln \overline{c_0} + \beta \ln \overline{c_1} + \beta^2\mu - \frac{1}{2}\frac{\beta^3}{\theta}\sigma^2 \\ V_0(0_L) &= \ln \overline{c_0} - \beta\theta \ln E[e^{-\frac{1}{\theta}\ln V_1(s_1)}] \\ &= \ln \overline{c_0} + \beta V_1(1) \\ &= \ln \overline{c_0} + \beta \ln \overline{c_1} + \beta^2\mu - \frac{1}{2}\frac{\beta^2}{\theta}\sigma^2 \end{aligned}$$

In period -1:

$$V_{-1} = \ln \overline{c_{-1}} - \beta\theta \ln E[e^{-\frac{1}{\theta}V_0(s_0)}]$$

B.3 SDF calculation

Pricing period 2 cash flow:

$$\begin{aligned} SDF(1_U, s_2) &= \beta \frac{e^{-\frac{1}{\theta}\ln c_2}}{E[e^{-\frac{1}{\theta}\ln c_2}]}(\frac{c_2}{\overline{c_1}})^{-1} = \beta(\frac{c_2}{\overline{c_1}})^{-1} \\ SDF(1_D, s_2) &= \beta \frac{e^{-\frac{1}{\theta}\ln c_2}}{E[e^{-\frac{1}{\theta}\ln c_2}]}(\frac{c_2}{\overline{c_1}})^{-1} = \beta(\frac{c_2}{\overline{c_1}})^{-1} \\ SDF(1, s_2) &= \beta \frac{e^{-\frac{1}{\theta}\ln c_2}}{E[e^{-\frac{1}{\theta}\ln c_2}]}(\frac{c_2}{\overline{c_1}})^{-1} \end{aligned}$$

Pricing period 1 cash flow:

$$
\begin{aligned}
SDF(0_E, s_1) &= \beta \frac{e^{-\frac{1}{\theta}V_1(s_1)}}{E[e^{-\frac{1}{\theta}V_1(s_1)}]}(\frac{\overline{c_1}}{\overline{c_0}})^{-1} \\
&= \beta \frac{e^{-\frac{1}{\theta}\beta \ln c_2}}{E[e^{-\frac{1}{\theta}\beta \ln c_2}]}(\frac{\overline{c_1}}{\overline{c_0}})^{-1} \\
SDF(0_L, 1) &= \beta(\frac{\overline{c_1}}{\overline{c_0}})^{-1}
\end{aligned}
$$

The risk-free rate from 0 to 1: $R_f(0_E) = \frac{1}{E[SDF(0_E, s_1)]} = \frac{1}{\beta}(\frac{\overline{c_1}}{\overline{c_0}})$, $R_f(0_L) = \frac{1}{\beta}(\frac{\overline{c_1}}{\overline{c_0}})$.

Pricing period 0 cash flow:

$$
SDF(-1, s_0) = \beta \frac{e^{-\frac{1}{\theta}V_0(s_0)}}{E[e^{-\frac{1}{\theta}V_0(s_0)}]}(\frac{\overline{c_0}}{\overline{c_{-1}}})^{-1}
$$

The risk-free rate from -1 to 0: $R_f(-1) = \frac{1}{E[SDF(-1, s_0)]} = \frac{1}{\beta}(\frac{\overline{c_0}}{\overline{c_{-1}}})$

B.4 Asset pricing

Consider an asset that pays $C_2^{1+\phi}$ in the period 2, where $\phi > 0$. We compute its price throughout the tree.

In period 1:

$$
\begin{aligned}
P_1(s_1) &= E[SDF(s_1, s_2)c_2^{1+\phi}] \\
&= \beta(\frac{c_2}{c_1})^{-1}c_2^{1+\phi} = \beta\overline{c_1}c_2^{\phi} \\
P_1(1) &= E[SDF(1, s_2)c_2^{1+\phi}] \\
&= \beta\overline{c_1}e^{\phi\mu + \frac{1}{2}(\phi^2 - 2\frac{\phi}{\theta})\sigma^2}
\end{aligned}
$$

In period 0:

$$
\begin{aligned}
P_0(0_E) &= E[SDF(0_E, s_1)P_1(s_1)] \\
&= \beta^2\overline{c_0}e^{\phi\mu+\frac{1}{2}(\phi^2-2\frac{\phi\beta}{\theta})\sigma^2} \\
P_0(0_L) &= SDF(0_L, 1)P_1(1) \\
&= \beta^2\overline{c_0}e^{\phi\mu+\frac{1}{2}(\phi^2-2\frac{\phi}{\theta})\sigma^2}
\end{aligned}
$$

B.5 Announcement premium

Let's focus on the returns from 0_E to 1_E. That is the announcement returns.

$$
\begin{aligned}
R_A = \frac{P_1(s_1)}{P_0(0_E)} &= \frac{\beta\overline{c_1}c_2^{\phi}}{\beta^2\overline{c_0}e^{\phi\mu+\frac{1}{2}(\phi^2-2\frac{\phi\beta}{\theta})\sigma^2}} \\
&= \frac{1}{\beta}\frac{\overline{c_1}}{\overline{c_0}}\frac{c_2^{\phi}}{e^{\phi\mu+\frac{1}{2}(\phi^2-2\frac{\phi\beta}{\theta})\sigma^2}} \\
E[R_A] = & \quad \frac{1}{\beta}\frac{\overline{c_1}}{\overline{c_0}}e^{\frac{\phi\beta}{\theta}\sigma^2}
\end{aligned}
$$

Note that $R_f(0_E) = \frac{1}{\beta}\frac{\overline{c_1}}{\overline{c_0}}$. Therefore $\frac{E[R_A]}{R_f(0_E)} = e^{\frac{\beta\phi}{\theta}\sigma^2} > 1$ as long as $\theta > 0$.

B.6 Implied variance

Consider the claim to the variance of the return from the end of period 0 to the end of period 1. Since the payoff is known and paid out at the end of period 0, we have:

$$
\begin{aligned}
IV_{0\to 1}(0_E) &= Var[\ln\frac{P_1(s_1)}{P_0(0_E)}] \\
IV_{0\to 1}(0_L) &= Var[\ln\frac{P_1(1)}{P_0(0_L)}]
\end{aligned}
$$

Clearly,

$$
\begin{aligned}
IV_{0\to 1}(0_E) &= \phi^2\sigma^2 \\
IV_{0\to 1}(0_L) &= 0
\end{aligned}
$$

B.7 PER Premium

Consider the price of $IV_{0\to1}(s_0)$ at time -1. It is

$$E[SDF(-1,s_0)IV_{0\to1}(s_0)] = \beta(\frac{\overline{c_0}}{\overline{c_{-1}}})^{-1}\frac{1}{2}\frac{e^{-\frac{1}{\theta}V_0(0_E)}\phi^2\sigma^2}{E[e^{-\frac{1}{\theta}V_0(s_0)}]}$$

The expected return of this asset is therefore

$$E[R_{Var}(s_{-1},s_0)] = \frac{\frac{1}{2}\phi^2\sigma^2}{IV_{0\to1}(-1)} = \frac{1}{\beta}(\frac{\overline{c_0}}{\overline{c_{-1}}})\frac{E[e^{-\frac{1}{\theta}V_0(0_E)}]}{e^{-\frac{1}{\theta}V_0(0_E)}}$$

Therefore:

$$\begin{aligned}\frac{E[R_{Var}(s_{-1},s_0)]}{R_f(-1)} &= \frac{E[e^{-\frac{1}{\theta}V_0(0_E)}]}{e^{-\frac{1}{\theta}V_0(0_E)}}\\ &= \frac{\frac{1}{2}e^{\frac{\beta^3}{2\theta^2}\sigma^2}+\frac{1}{2}e^{\frac{\beta^2}{2\theta^2}\sigma^2}}{e^{\frac{\beta^3}{2\theta^2}\sigma^2}} > 1 \text{ iff } \beta < 1\end{aligned}$$

C Data Appendix

Our VIX data come from CBOE's website, and option data OptionMetrics. While the VIX data are straightforward to use, the handling of the option data is more involved. Below we describe our data construction process in detail.

Starting with a big panel of option prices, we first get the data to underlying-expiration-strike price-put/call-day level, i.e. for a put or call option on a certain underlying that has a certain expiration date and strike price we should have one price per day. There are some cases where there are two prices per day.[13] In those cases, we take the average of those two available prices.

Having a panel at the underlying-expiration-strike price-put/call-day level, we take the average of bid and ask to get the price of an option. This price can be missing, however, even for large underlying such as the S&P 500 index. This is because price inquiries can be rare for deeply in-the-money or out-of-money options. In the event that a price becomes missing and reappear in a future date, we forward fill the price, assuming a return of zero. If the price

[13] Such cases are because there are two types of options, e.g. standard monthly options and weekly options, that happen share the same underlying, expiration, strike price, and put/call and are both outstanding on the same day.

becomes missing forever, we replace the first missing price with zero if the option is a call and the last available call price is less than the put price of the same strike and expiration, and with the last available price if the last available call price is greater than or equal to the put price. Similarly, if the option is a put, we replace the missing price with 0 if its price is less than the call with the same strike and expiration, and with the last available price if it is the greater than the put price. This logic is to roughly impute a zero final return if the option is in the money, and a return of -100% if it is out of money.[14] While this operation is conceptually importantly, our results are robust to alternative imputation methods such as assuming all final returns are 0.

Having non-missing prices, we construct the synthetic variance claims using these S&P 500 options and compute their returns. We construct these variance claims following the formulas in Bakshi et. al. (2003), with additional data cleaning procedures taken from the construction of the VIX index, which are documented on the VIX white paper, available on CBOE's website. We also describe our methodology in detail below.

Overall, the portfolio on any given day consists of out-of-money options, which are call options with strike prices higher than the previous close price of the underlying, and put options with strike prices lower than that close price. Out-of-money options with zero bid prices are excluded from the portfolio. Also, those with two consecutive zero bids between them and the at-the-money strike price are also excluded. For instance, suppose a call with strike 100 has a non-zero bid price, and on that day the at-the-money strike price is 30. Let's say the two strike prices immediately lower than 100 is 95 and 90, and calls with those two strikes prices both have zero bids. Then the call option with strike price of 100 will be excluded even though it is an out-of-money option with a non-zero bid. These data exclusion logic is adopted from the CBOE's methodology in constructing the VIX index. Additionally, we require that an equal number of puts and calls are included in our variance-paying portfolio. This is to make sure that the portfolios are balanced between puts and calls and do not have strong directional delta exposure.

Having the sample we now discuss the weight of each option in the portfolio of variance swap. Say an option has a strike price of K, and the two nearby strike prices flanking K for that underlying-expiration-day are K^- and K^+. Let the underlying's close price on the previous trading day be S. For the second moment portfolio, the relative weight on the option with strike K is $\frac{(K^+ - K^-)}{2}\frac{1-\log(K/S)}{K^2}$. If the strike price is the highest or the lowest

[14] In the context of individual stock options, this logic is expensive due to the size of the data. We instead replace all final missing price with last day's price, and additionally verify that our results do not change appreciably if we replace all final missing prices with 0, of if we conditionally replace all missing prices with 0 or last day's price based on whether last day's price is greater than a dollar.

for that underlying-expiration-day, the weight is then $\frac{(K-K^{-})}{2}\frac{1-\log(K/S)}{K^2}$ or $\frac{(K^{+}-K)}{2}\frac{1-\log(K/S)}{K^2}$, respectively. We then rescale these relative weights so that they add up to 1 for each underlying-expiration-day. Weighted-returns on these portfolios are then computed. In the context of S&P 500 option portfolios, these returns are used as is because the data can be manually examined to make sure that they are free of influential data errors. For individual stock options such manual examination is not possible. We instead winsorize these returns at the 0.5 and 99.5 percentiles, and additionally verify that our results are robust to the chosen percentiles.